Diskussionsschriften

What’s Really the Story with this Balassa-Samuelson Effect in the CEECs?

Martin Wagner
Jaroslava Hlouskova

04-16

September 2004

Universität Bern
Volkswirtschaftliches Institut
Gesellschaftstrasse 49
3012 Bern, Switzerland
Tel: 41 (0)31 631 45 06
Web: www.vwi.unibe.ch

What's Really the Story with this Balassa-Samuelson Effect in the CEECs?*

Martin Wagner †
University of Bern
Department of Economics

Jaroslava Hlouskova
Institute for Advanced Studies Vienna
Department of Economics and Finance

September 23, 2004

Abstract

This paper offers a detailed assessment of the Balassa-Samuelson (BS) effect in eight Central and Eastern European countries (CEEC8). Several features distinguish this study from others: First, we investigate a variety of specifications of *extended* models. Non-homogeneity of wages, deviations from PPP in tradables and demand side variables are found to importantly contribute to explain inflation differentials. Second, a variety of specifications is investigated. Third, we rely upon bootstrap inference for panel unit root and panel cointegration analysis. The bootstrap results are rather clear: No evidence for cointegration remains when resorting to bootstrap inference. To quantify the *bias* that may arise from incorrectly using cointegration techniques, we also quantify the BS effect from equations containing (nonstationary) 'cointegration' terms. Fourth, we present inflation simulations based on well specified scenarios.

The results are as follows: Evidence for the BS effect is found. The BS effect is, however, rather small (around half a percent per annum) and not sufficient to explain the observed inflation differentials between the CEEC8 and the EU11. Using, despite the lacking evidence, cointegration techniques results throughout in substantially larger estimated effects. This suggests that studies relying upon cointegration may have overestimated the BS effect.

The additional explanatory variables in the *extended* BS models allow for a satisfactory modelling of the observed inflation rates. The mean inflation simulations for the CEEC8 countries, based on the extended models, range from 2.77% for the Slovak Republic to 6.75% for Poland. These are well above the 2% inflation objective for the European Monetary Union.

JEL Classification: F02, O40, O57, P21, P27

Keywords: Balassa-Samuelson effect, Central and Eastern Europe, transition economies, non-stationary panels, bootstrapping, inflation simulations

---

*Financial support from the Jubiläumsfonds of the Österreichische Nationalbank under grant Nr. 9557 is gratefully acknowledged. Thanks to M. Kotov for assisting in the data collection and management. We would furthermore like to thank Rumen Dobrinsky for sharing some of his data with us.

†Corresponding author: Gesellschaftsstrasse 49, CH-3012 Bern, Switzerland; Tel. ++41 31 631 4778, Fax: ++41 31 631 3992, email: `martin.wagner@vwi.unibe.ch`. Part of this work has been done whilst visiting the Economics Department of Princeton University, whose hospitality is gratefully acknowledged.

1 Introduction

In this paper we present a detailed econometric study of the Baumol-Bowen (BB) and Balassa-Samuelson (BS) effect for eight Central and Eastern European Countries (CEECs). Studies of these effects have been offered in abundance in recent years. In particular the recent EU enlargement and the subsequent entry of the new member states into the European Monetary Union spur the interest in studies of the real exchange rate and inflation behavior of these economies. Different *structural inflation rates* across monetary union member states may pose a challenge for common monetary policy, see e.g. Sinn and Reutter (2001). Recent interest in the BS model stems from the fact that it explains differences in inflation rates (respectively real exchange rates) by different productivity growth differentials between the tradables and non-tradables sectors across countries, see the discussion of the model in Section 2. Since larger productivity differentials are often observed in catching-up economies the BS model has been prominent in explaining higher inflation rates in, respectively real exchange rate appreciations of, catching-up economies, see Canozoneri et al. (1999) for OECD country evidence or Mihaljek and Klau (2004) for a study on CEECs.

In our study we try to improve over current practice in the empirical BS literature in several directions. First, the highly stylized theoretical model rests upon a variety of assumptions that lead to a purely supply side based explanation of real exchange rates respectively inflation rates. We check for the presence of demand side effects and find in particular real per capita GDP important. This finding is consistent with the extension of the BS model presented in Bergstrand (1991). Furthermore we assess in detail the validity of two additional key assumptions of the BS model. These are wage homogeneity across sectors and the prevalence of purchasing power parity (PPP) in tradables. Our econometric analysis leads us to refute both. We thus work with the so called *extended* versions of the models, that relax these two assumptions. Also, we specify a multitude of equations based on the model. We differentiate the estimation equations along two dimensions. The first is the choice of the dependent variable. Since the theoretical model is specified for a two-sector economy, composed of tradables and non-tradables, we use in the *narrow* specifications only the prices in these two sectors, respectively the real exchange rate with respect to these two sectors as dependent variables. The *broader* specifications are less theory driven and use the GDP deflators respectively the corresponding real exchange rates as dependent variables. The equations with the narrow dependent variables in general show better fit, as expected. The second dimension along which we distinguish the equations is the choice of the BS variable, see Section 2 for details. The five choices concerning the BS variable differ e.g. with respect to how sectoral wages are considered. These two specification issues, narrow and wide measure of dependent variables and choices of BS variables, have not yet been treated systematically in the literature.

Second, we acknowledge in our econometric analysis the fact that econometric methods for small nonstationary panels are known to behave unsatisfactorily. Especially panel unit root and panel cointegration tests are known to suffer from severe distortions in small samples, see Hlouskova and Wagner (2004a) or Gutierrez (2003). We try to overcome these limitations by resorting to bootstrapping methods. Various bootstrap algorithms are implemented and lead to similar results: Unit root nonstationarity is found to be widespread amongst the variables. However, essentially no evidence for cointegration is found, when resorting to bootstrap inference. This finding stands in stark contrast with other studies that rely upon panel cointegration methods, see e.g. Egert (2002) or Egert et al. (2002). To assess the *bias* that is introduced by incorrectly resorting to cointegration techniques, we also quantify the BS effect for well specified equations including error correction terms. We term an equation well specified if all coefficient signs are in line with theory, this includes the coefficient in the 'cointegrating' relationship. The results are quite clear for our data for all equations: Using cointegration leads to an over-estimation of the BS effect throughout, partly substantially (by a factor up to four).

We find ample evidence for the BS effect being present. With an average value of about half a percent per annum, it is however too small to explain observed inflation differentials between the CEEC8 and the EU11.[1] This finding is consistent with the above mentioned observation that several key assumptions of the standard BS model are not supported by the data. Thus, the pure BS effect alone cannot be expected to be too powerful in explaining inflation differentials, respectively real exchange rate movements. It is the inclusion of variables like deviation from PPP in tradables, relative sectoral wages, real per capita GDP or total consumption that allow for well specified BS type equations with good fit. We therefore base our inflation simulations not just upon the estimated BS effects, but include also the other

[1] The 'foreign country' used in our study, denoted by EU11, is the aggregate of eleven incumbent EU member states: Austria, Belgium, Denmark, Finland, France, Germany, Great Britain, Italy, The Netherlands, Spain and Sweden. The other incumbent EU member states are omitted because of lacking sectoral data. The sample period for the empirical analysis is 1993–2001 with annual data.

explanatory variables in our inflation simulations, see the details, in particular also concerning the scenario assumptions, in Section 7. The bottom line of the large set of results can be roughly summarized as follows: The mean inflation projection is between 2.77% for the Slovak Republic to 6.75% for Poland. The mean prediction for the aggregate inflation of the CEEC8 is 5.43%, with a standard deviation over specifications of about 1.2% inflation rate. These numbers, are well above the inflation objective of 2% formulated for the European Monetary Union. Also the fact that relatively large inflation differences are predicted across countries, may pose a challenge for monetary policy in an enlarged monetary union.

The paper is organized as follows. In Section 2 we start with a discussion of the theoretical model and the relationships derived thereof for the econometric analysis. Section 3 is devoted to a description of the data and a preliminary graphical investigation of some key elements of the BS model. In Section 4 panel unit root tests are performed and Section 5 is devoted to panel cointegration analysis. In Section 6, based on the results of the previous sections, appropriate equations are specified and the BB and BS effects are quantified. In Section 7 we discuss and present the inflation simulations and Section 8 briefly summarizes and concludes. Three appendices follow the main text: Appendix A contains a detailed description of the data, their sources and preliminary variable transformations. In Appendix B a multitude of additional empirical results is collected and in Appendix C the implemented bootstrap algorithms are briefly described.

2 The Baumol-Bowen and the Balassa-Samuelson Effect

Balassa (1964) and Samuelson (1964) present models in which *different* productivity growth differentials between the *tradables* and *non-tradables* goods sectors *across countries* are an important factor in explaining real exchange rate movements, respectively differences in the evolution of national price levels.[2]

The model is formulated in terms of a two-sector small open economy. The small open economy assumption implies that the world interest rate R and the world market price of tradables P^T are taken as given. Both sectors, tradables (T) and non-tradables (N), are described by their sectoral production functions, which are for algebraic simplicity assumed

[2]Recently Ghironi and Melitz (2003) presented a very interesting stochastic general equilibrium model with heterogeneous firms that leads to BS type effects. An econometric analysis of that model will be an interesting challenge for the BS community. An earlier general equilibrium analysis of the BS model is given by Asea and Mendoza (1994).

to be Cobb-Douglas:[3]

$$\begin{aligned} Y^T &= A^T (K^T)^{1-\alpha^T} (L^T)^{\alpha^T} \\ Y^N &= A^N (K^N)^{1-\alpha^N} (L^N)^{\alpha^N} \end{aligned} \tag{1}$$

where Y^s, with $s \in \{T, N\}$, denotes real output in sector s; A^s, K^s and L^s denote (total factor) productivity, capital and labor in the respective sector; and α^s denotes labor intensity in each sector. The productivities and the labor intensities are allowed to differ across the two sectors. Both sectors are assumed to be composed of perfectly competitive firms and production factors are assumed to be fully utilized. The assumptions imply that only the supply side of the economy influences the evolution of the real exchange rate. The potential effect of demand side factors for the evolution of the real exchange rate in the CEECs is tested in Section 6.

The assumption of perfect competition in both sectors leads to the following first order conditions for profit maximization, with W^T and W^N denoting the wages in the tradables and non-tradables sector.[4]

$$\begin{aligned} R &= (1-\alpha^T) A^T \left(\frac{L^T}{K^T}\right)^{\alpha^T} \\ &= P^{rel} (1-\alpha^N) A^N \left(\frac{L^N}{K^N}\right)^{\alpha^N} \\ W^T &= \alpha^T A^T \left(\frac{L^T}{K^T}\right)^{-(1-\alpha^T)} \\ W^N &= P^{rel} \alpha^N A^N \left(\frac{L^N}{K^N}\right)^{-(1-\alpha^N)} \end{aligned} \tag{2}$$

where $P^{rel} = P^N/P^T$ denotes the relative price of non-tradables.

Concerning the labor market, in the *standard* Balassa-Samuelson model perfect labor mobility across the two sectors is assumed. This results in wage homogeneity, $W^T = W^N$. Under this additional assumption the above equations can be solved to obtain the following expression for the logarithm of relative prices.[5]

$$p^{rel} = c + \frac{\alpha^N}{\alpha^T} a^T - a^N \tag{3}$$

where c is a constant depending upon the exogenously given factor intensities (α^T, α^N) and the interest rate. Throughout the letter c is used to denote constants in the various equations, those are not necessarily the same across equations. The above equation (3) displays the

[3]The choice of Cobb-Douglas functions, with its algebraic convenience of leading to simple log-linear equilibrium relationships, is of course an approximation. Thus, some flexibility in the empirical modelling might be required.

[4]Throughout the discussion we consider the tradables good as the numeraire.

[5]Lower case letters indicate logarithms of variables throughout.

link between the relative prices in the two sectors and, up to the factor $\frac{\alpha^N}{\alpha^T}$, the relative productivities. This effect is known in the literature as the *Baumol-Bowen* effect, see Baumol and Bowen (1966). The underlying logic of the argument is simple: For simplicity of the verbal argument assume for the moment that $\alpha^T = \alpha^N$. Assume further that productivity growth is faster in the tradables sector than in the non-tradables sector, i.e. $\Delta a^T > \Delta a^N$. Now, if productivity grows faster in the tradables sector, this allows for wages to grow faster in this sector (given the exogenous world market prices for tradables and capital). Due to the assumed labor mobility, the non-tradables sector has to pay the same wages as the tradables sector. This implies, due to lower productivity growth, that the non-tradables sector has to raise its prices (faster) in order to remain profitable. Thus, higher productivity growth in the tradables sector leads to higher inflation in the non-tradables sector. Note that in many countries the labor intensity is higher in the non-tradables sector than in the tradables sector, i.e. $\alpha^N > \alpha^T$, which reinforces the above argument where we assumed identical intensities for simplicity.

Surprisingly, many empirical studies like Alberola and Tyrväinen (1998), Coricelli and Jazbec (2004a), Coricelli and Jazbec (2004b), Halpern and Wyplosz (2002) or Sinn and Reutter (2001) that claim to study the Balassa-Samuelson effect are in fact studying the Baumol-Bowen effect. The imprecision in the distinction may stem from the fact that the relative price of non-tradables to tradables is often used as an *internal* measure for the real exchange rate. This measure, however, will in general differ substantially from other real exchange rate variables, based on the GDP or CPI deflators or also the trade weighted real exchange rate. Note also that the Baumol-Bowen effect is only concerned with domestic variables, thus in particular it cannot explain any *inflation differentials across* countries. The Baumol-Bowen effect is only one important part of the Balassa-Samuelson effect, as will become clear below.

Without the assumption of sectoral labor mobility and the implied wage homogeneity, the above equation (3) is modified to

$$p^{rel} = c + \frac{\alpha^N}{\alpha^T} a^T - a^N + \alpha^N (w^N - w^T) \tag{4}$$

The interpretation of this *extended* Baumol-Bowen effect is similar to the explanation given above. Now, for example, lower wage growth in the non-tradables sector can mitigate the relative inflation pressure.

The Balassa-Samuelson effect itself combines the above domestic Baumol-Bowen effect with the (evolution of the) real exchange rate. Starred variables henceforth denote the foreign country, or the rest of the world. In our empirical analysis the foreign country is given by, as already stated, the EU11. The real exchange rate for a country is defined as $Q = \frac{EP^*}{P}$, where E denotes the nominal exchange rate (local currency per Euro) and P and P^* denote the domestic and foreign aggregate price levels. Throughout the paper variables for the EU11 are indicated with a '*'. The aggregate price levels are weighted averages (weighted by expenditure shares δ) of the sectoral price levels, i.e. in logarithms they are given by:

$$\begin{aligned} p &= (1-\delta)p^T + \delta p^N \\ p^* &= (1-\delta^*)p^{T*} + \delta^* p^{N*} \end{aligned} \tag{5}$$

Combining the above price level decompositions with the definition of the real exchange rate directly leads to

$$q = (e + p^{T*} - p^T) - \delta(p^N - p^T) + \delta^*(p^{N*} - p^{T*}) \tag{6}$$

Thus, the (logarithm of the) real exchange rate is seen to depend upon three factors: The first is the real exchange rate in the tradables sector. It is commonly assumed that PPP holds in the tradables sector, this implies $e + p^{T*} - p^T = 0$. Thus, under this assumption the first term vanishes. The second and third term are the relative prices of non-tradables in both countries, weighted by their shares in the overall price level. Inserting the expressions for the relative prices found above, leads to the Balassa-Samuelson model, that explains the real exchange rate in terms of productivity differentials at home and abroad

$$q = c + (e + p^{T*} - p^T) - \delta\left(\frac{\alpha^N}{\alpha^T}a^T - a^N\right) + \delta^*\left(\frac{\alpha^{N*}}{\alpha^{T*}}a^{T*} - a^{N*}\right) \tag{7}$$

Given that PPP holds in the tradables sector, the real exchange rate is given by:

$$q = c - \delta\left(\frac{\alpha^N}{\alpha^T}a^T - a^N\right) + \delta^*\left(\frac{\alpha^{N*}}{\alpha^{T*}}a^{T*} - a^{N*}\right) \tag{8}$$

The above equation (8) implies, for sufficiently similar labor intensities and expenditure shares, that the real exchange rate of the country appreciates ($\Delta q < 0$), if its sectoral productivity growth rate differential is larger than the productivity growth differential abroad. The fact that this differential is often found to be bigger in faster growing or catching-up economies, makes the Balassa-Samuelson model a widely used model for explaining real exchange rate appreciations. Employing once again the definition of the real exchange rate, the above equation (8) can be modified and differenced to describe inflation differentials across countries:

$$\Delta p - \Delta p^* = c + \Delta e + \delta\left(\frac{\alpha^N}{\alpha^T}\Delta a^T - \Delta a^N\right) - \delta^*\left(\frac{\alpha^{N*}}{\alpha^{T*}}\Delta a^{T*} - \Delta a^{N*}\right) \tag{9}$$

The inflation differential depends upon nominal exchange rate movements and the differences in the sectoral productivity growth differentials across countries. In a monetary union, the nominal exchange rate is by construction fixed, and inflation differentials are, according to the model, solely determined by productivity growth differentials across member states of a monetary union.

As for the Baumol-Bowen effect discussed above, also for the Balassa-Samuelson effect the assumption of wage homogeneity across sectors can be relaxed. This results in the following generalization of equation (9), now again in levels:

$$\begin{aligned} p - p^* \quad &= \quad c + e + \delta\left(\tfrac{\alpha^N}{\alpha^T}a^T - a^N + \alpha^N(w^N - w^T)\right) \\ &\quad -\delta^*\left(\tfrac{\alpha^{N*}}{\alpha^{T*}}a^{T*} - a^{N*} + \alpha^{N*}(w^{N*} - w^{T*})\right) \end{aligned} \tag{10}$$

Abstaining from the assumption of PPP for traded goods, we obtain the following reformulation of equation (9)

$$\Delta p - \Delta p^* \quad = \quad c + \Delta p^T - \Delta p^{T*} + \delta\left(\frac{\alpha^N}{\alpha^T}\Delta a^T - \Delta a^N\right) - \delta^*\left(\frac{\alpha^{N*}}{\alpha^{T*}}\Delta a^{T*} - \Delta a^{N*}\right) \tag{11}$$

which holds without any assumption on the nominal exchange rate. Now inflation differentials depend upon tradables inflation differentials and the differences in productivity growth differentials. Of course, as above, also the extension allowing for non-homogenous wages can (and will) be investigated:

$$\begin{aligned} \Delta p - \Delta p^* \quad &= \quad c + \Delta p^T - \Delta p^{T*} + \delta\left(\tfrac{\alpha^N}{\alpha^T}\Delta a^T - \Delta a^N + \alpha^N(w^N - w^T)\right) - \\ &\quad \delta^*\left(\tfrac{\alpha^{N*}}{\alpha^{T*}}\Delta a^{T*} - \Delta a^{N*} + \alpha^{N*}(w^{N*} - w^{T*})\right) \end{aligned} \tag{12}$$

From the above relationships various variables that correspond to the Balassa-Samuelson effect can be derived. The variable $BS_{it} = \delta_{it}a_{it}^{rel} - \delta_t^* a_t^{rel*}$ follows from equation (8) after setting $\alpha^N = \alpha^T$ in both the CEE country and the EU11, with $a^{rel} = a^T - a^N$.[6] Here and throughout the paper in the double sub-script it, i is the country and t the time index. These are dropped when unnecessary. The shares δ_{it} can be easily computed by $\delta_{it} = \frac{Y_{it}^N}{Y_{it}^T + Y_{it}^N}$. Taking into account the non-homogeneity of wages (established below), the variable $BSE1_{it}$ is

[6]We furthermore experimented with variables that contain $\frac{\alpha^N}{\alpha^T}a^T - a^N$ instead of $a^T - a^N$. These variables, despite their theoretical appeal do not lead to satisfactory econometric analysis and results. This may inter alia reflect that the sectoral production functions are not exactly Cobb-Douglas.

computed as follows $BSE1_{it} = \delta_{it}(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - \delta_t^*(a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})$, with $w^{rel} = w^N - w^T$. Implicitly setting $\delta_{it} = \delta_t^*$, i.e. ignoring differences in the sectoral composition across the CEE countries and the EU11, defines the variable $BSE2_{it} = (a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - (a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})$. Finally also the differential of relative productivities, $a_{it}^{rel} - a_t^{rel*}$, is used as a BS variable. This latter choice is probably the most widely used variable, despite or because of neglecting some Cobb-Douglas related constants.

As indicated in the introduction in the following sections the above relationships are investigated using panel unit root and panel cointegration techniques. Usage of this type of techniques rests upon the first tested assumption of unit-root non-stationarity for the macro-variables used. Unit-root non-stationarity combined with the presence of cointegration as laid out by the above relationships leads to error-correction models for the evolution of the (rate of appreciation of the) real exchange rate, respectively of the inflation differentials. In the empirical analysis we furthermore address the potential impact of demand side factors on the evolution of prices and exchange rates.

3 Data and Preliminary Investigations

The study is conducted for eight Central and Eastern European countries (CEEC8): the Czech Republic (CZE), Estonia (EST), Hungary (HUN), Latvia (LVA), Lithuania (LTU), Poland (POL), the Slovak Republic (SVK) and Slovenia (SVN). The foreign country in the empirical study is, as mentioned, comprised by the aggregate of eleven incumbent EU (EU11) member states, these are the EU15 excluding Greece, Ireland, Luxembourg and Portugal. These four countries are omitted because of incomplete data. Note, however, that these are all relatively small economies that are rather unrelated to our CEE countries. Thus, the effect of the omission of these countries in the construction of the foreign country can be expected to be modest. The data are annual and the sample period is 1993–2001. This is also the sample period used throughout the econometrics in the subsequent sections.

The first decision to make is, of course, the sectoral classification. We decide to take NACE sectors C (mining and quarrying), D (manufacturing) and E (electricity, gas and water supply) as our tradables (T) sector. Non-tradables (N) is composed of NACE sectors F (construction) to K (real estate and business activities). NACE sectors A and B are aggregated to agriculture (AGR) and sectors L to P are aggregated to the public sector (PUB). See Table 20 in Appendix A for details.

A description of all available variables and their sources is given in Tables 21 and 22. For reference purposes all variable transformations prior to econometric analysis are collected in Table 23. The precise construction of the EU11 aggregates for the tradable and the non-tradable sectors is contained in Table 24. These are aggregated using sectoral output weights. All these tables describing the data and preliminary variable transformations are contained in Appendix A.

With the chosen classification, about 70 to 80 % of the economy are taken into account, see the right block of Table 1. The two neglected sectors, agriculture and the public sector, have quite substantial inflation rates, see columns five and six of Table 1. For this reason, we have decided to specify the empirical analogues of equations (7) to (12), derived in the previous section, with two different price indices respectively two different real exchange rate measures.[7] The two price differentials are given by $p_{it}^{GDP} - p_t^{GDP*}$, i.e. the difference in the (logarithms of the) GDP deflators. Following Harberger (2004) and based on the fact that our model is specifying the supply side of the economy, we have decided to use the GDP deflator as our broad price measure. The other possible choice for a broader price aggregate would be the *consumer price index* (CPI). The correlation between the GDP based inflation rates and the HICP (Harmonized Index of Consumer Prices) inflation rates is close to one for most countries. Thus, no qualitative differences in the results have to be expected.[8] The second price differential chosen is given by $p_{it}^{T+N} - p_t^{(T+N)*}$, i.e. by the differential of the log price levels only in the two sectors tradables and non-tradables. Similarly to the two price variables also the corresponding two real exchange rate measures have been chosen, $q_{it} = e_{it} + p_t^{GDP*} - p_{it}^{GDP}$ and $q_{2,it} = e_{it} + p_t^{(T+N)*} - p_{it}^{T+N}$. From the definition of the variables it immediately follows that the predictions concerning the coefficient signs in the p-equations are opposite those for the q-equations. Specifying two sets of equations, based on a narrow and a wide price respectively real exchange rate measure, allows us to assess the effect of the choice of dependent variable on the results. This is an issue up to now entirely neglected in the empirical literature. In the sequel we denote with p-equations the equations with the two price (differentials) as dependent variables and with q-equations the equations

[7] The empirical specifications will partly include further explanatory variables. All equations include relative wage terms and terms related to the real exchange rate of tradables. See the discussion below.

[8] The correlation between GDP deflator and HICP inflation rates over the period 1994–2001 is e.g. 0.95 for Sweden or 0.86 for the Netherlands. The average correlation across the EU11 is 0.81. For only two countries is the correlation below 0.7, Belgium and Finland. Not only the correlation between the HICP and the GDP deflator inflation is very high, also the dynamics of the two variables are very similar for all countries.

Country	Δa^T	Δa^N	Δp^T	Δp^N	Δp^{AGR}	Δp^{PUB}	Sectoral shares in total output T	N	AGR	PUB
Averages over 1994–2001										
CZE	5.15	2.17	5.14	6.51	6.39	12.06	0.35	0.47	0.05	0.13
EST	6.38	5.72	12.53	14.27	11.68	18.43	0.24	0.50	0.08	0.18
HUN	6.87	0.92	11.77	15.70	11.68	15.01	0.28	0.46	0.06	0.20
LVA	6.92	6.06	6.01	10.94	4.38	17.08	0.28	0.47	0.09	0.17
LTU	6.07	2.02	12.84	14.78	8.71	23.52	0.26	0.47	0.12	0.15
POL	7.99	2.66	7.93	18.61	12.36	17.03	0.33	0.45	0.06	0.16
SVK	3.82	2.30	5.86	8.03	5.31	6.33	0.31	0.47	0.06	0.16
SVN	6.96	2.15	9.47	12.31	8.92	11.11	0.33	0.44	0.04	0.20
EU11	2.84	1.01	1.25	2.03	0.70	2.82	0.24	0.53	0.02	0.21
Averages over 2000–2001										
CZE	5.6	7.9	2.2	0.9	9.1	7.8	0.35	0.49	0.05	0.11
EST	8.4	7.1	5.3	5.5	8.1	4.1	0.24	0.52	0.06	0.17
HUN	4.8	1.6	7.4	9.5	6.7	10.9	0.31	0.45	0.05	0.19
LVA	4.8	7.3	1.2	3.1	7.4	6.5	0.26	0.50	0.08	0.16
LTU	12.7	5.8	5.5	1.3	-1.7	0.8	0.27	0.48	0.11	0.15
POL	6.8	3.1	1.3	8.0	14.0	16.1	0.34	0.46	0.05	0.15
SVK	0.3	1.6	4.2	7.6	9.5	0.0	0.31	0.47	0.05	0.17
SVN	6.6	1.7	6.4	11.8	9.3	11.2	0.33	0.43	0.04	0.20
EU11	2.3	0.8	1.8	2.2	3.2	3.2	0.23	0.54	0.02	0.21

Table 1: Sectoral productivity growth rates, sectoral inflation rates and sectoral output shares. The top panel displays the average annual growth rates over the period 1994–2001 and the lower panel over the period 2000–2001.

that have the real exchange rate measures as dependent variables.

In Table 1 we also display the average sectoral productivity growth rates in the tradables and non-tradables sectors. Consider the productivity growth rates first. For the larger period, displayed in the top panel, it holds in all CEECs and the EU11 that productivity grows faster in the tradables sector than in the non-tradables sector. For the shorter period 2000–2001 this does not hold for the Czech Republic, Latvia and the Slovak Republic. Also note that the differentials vary substantially across countries. For example in the Slovak Republic the productivity growth differential (over the period 1994–2001) is smaller (1.52%) than in the EU11 (1.83%). Thus, we can already expect substantial differences across countries, concerning the extent of dual inflation pressures via sectoral productivity differentials.

Concerning sectoral inflation rates, we see in columns three and four that (again for the longer period) the non-tradables sector has a higher inflation rate than the tradables sector. For the shorter period again some opposite inflation dynamics occur, in the Czech Republic and in Lithuania.

Summing up the information from Table 1, some key facts in line with the Baumol-Bowen model are present in the data for the CEECs: Higher productivity growth rates in the tradables sector and higher inflation rates in the non-tradables sector. In Figure 1 roughly the same information is also shown graphically. For all countries and the EU11 we display, over the period 1993–2001 in solid lines the relative price of non-tradables to tradables, in fine dashed lines the relative productivity of tradables to non-tradables and in dashed lines the relative wages in the non-tradables sector relative to wages in the tradables sector. Thus, the relative prices and relative productivities are displayed in such a way that they should grow over time, if behaving according to the model with higher productivity growth in tradables and higher inflation in non-tradables. *Wage homogeneity* across sectors implies that the relative wages should not exhibit trending behavior. The evidence is mixed. Concerning relative wages we observe stable relative wages in the Czech Republic, Poland, Slovenia and the EU11, in other countries wage homogeneity seems to be violated.[9] Relative prices and productivities exhibit co-movements, with differing degrees of synchronicity. E.g. in Lithuania there is an almost one-to-one relation between relative prices and productivities.

Table 1 and Figure 1 allow for a first graphical assessment of the prevalence of a Baumol-Bowen effect in the CEECs. In the following Figure 2 we take a first look at a potential Balassa-Samuelson effect in the CEECs with respect to the EU11. The figure displays for three different periods the *differential* of the relative productivity growth rates in the CEECs, Δa^{rel}, to the relative productivity growth rate in the EU11, Δa^{rel*} against the *inflation rate differential* between the CEEC countries, Δp^{T+N} and the EU11, $\Delta p^{(T+N)*}$.[10] In its standard version, the Balassa-Samuelson effect implies a positive correlation between sectoral productivity growth differentials and inflation differentials, compare e.g. equation (9). This is supported by Figure 2. The correlations are 0.458 for the longest period 1994–2001, 0.836 for the period 1996–2001 and 0.419 for the shorter period 2000–2001. We already know from the above discussion that over the period 2000–2001 in several countries behavior not supporting the standard BS setup has been observed. This translates into the lower correlation over that short period. One further important observation can be made in the figure. For three or

[9]Formal econometric tests for wage homogeneity will be presented in the following sections, in the context of panel unit root and panel cointegration analysis. The tests reject the null hypothesis of wage homogeneity in the panel of eight CEE countries.

[10]Note that the inflation rates are only computed for the tradables and non-tradables sectors. Similar pictures with the GDP deflators exhibit also positive correlation, albeit slightly less pronounced.

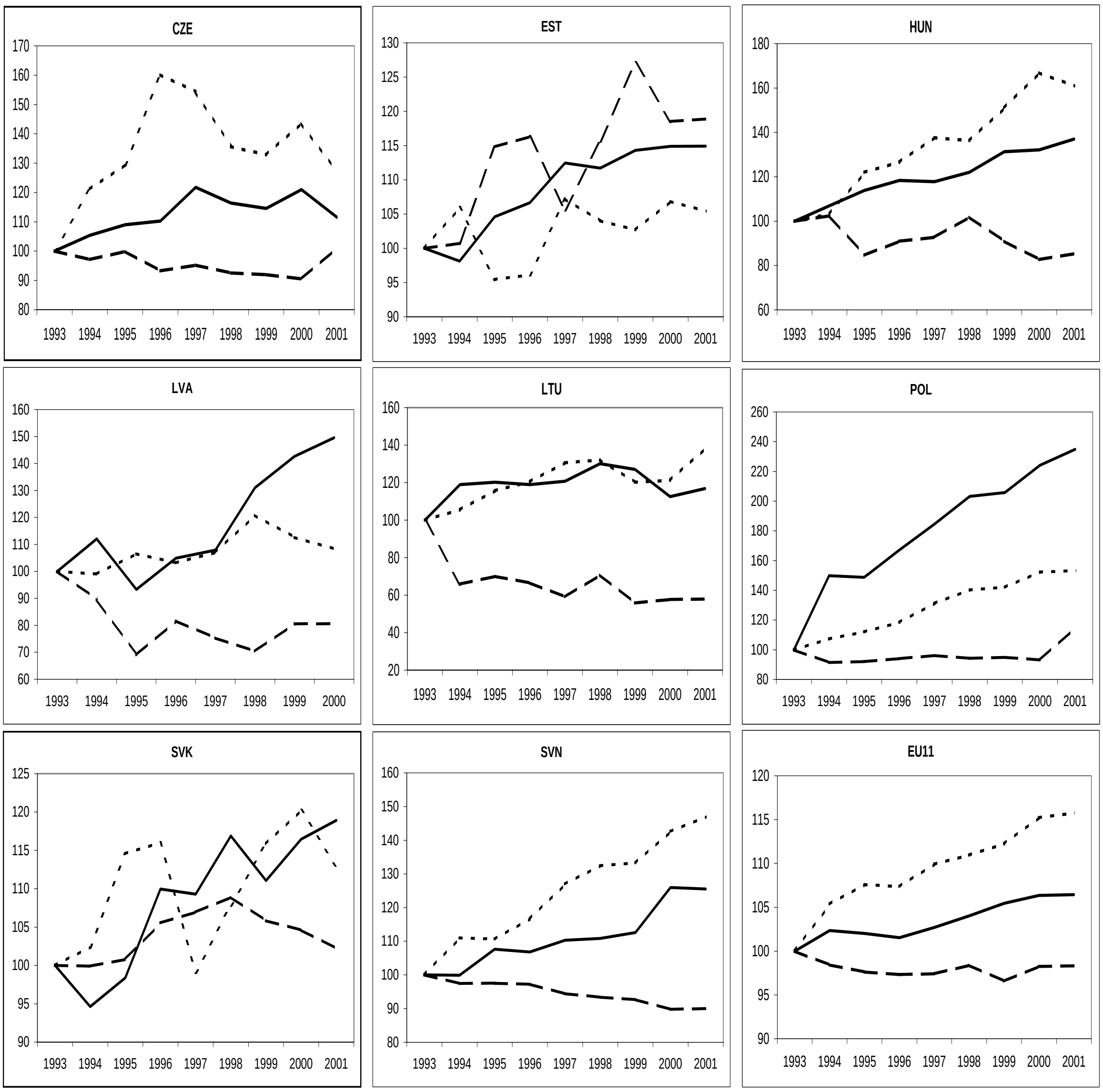

Figure 1: Solid lines: relative prices of non-tradables to tradables (N/T); fine dashed lines: relative productivities in tradables and non-tradables sector (T/N); dashed lines: relative wages in non-tradables and tradables sector (N/T). All quantities normalized to 100 in 1993.

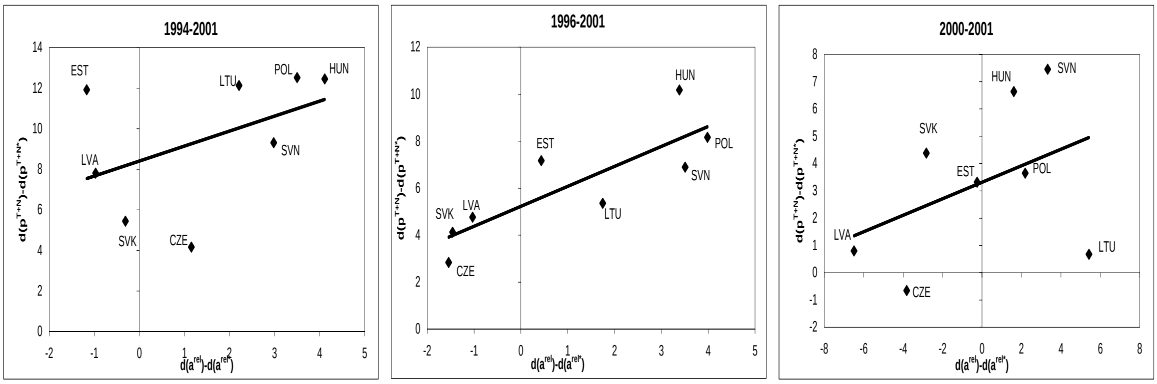

Figure 2: Relative productivity (T/N) and inflation differentials to the EU11. The inflation rates are computed only over the tradables and non-tradables sectors. The left chart displays the averages over the period 1994–2001, the chart in the middle displays the averages over the period 1996–2001 and the right chart over the period 2000–2001.

four out of eight countries, depending upon the period, the relative productivity differential growth rate is smaller than in the EU11. Thus, for these countries and these periods the standard BS model actually implies *smaller* inflation in the CEE countries than in the EU11. Combining this with the observed higher inflation rate in all CEE countries compared to the EU11 directly implies that the contribution of the BS term, which ever way measured, to inflation will be limited, despite the positive *unconditional* correlation. The model thus needs to be augmented by further explanatory variables, like the extensions discussed in Section 2 or demand side variables discussed as below in Section 6.

The evidence gained in this section by considering averages and also by graphical inspection of some key ratios and relationships is *grosso modo* sufficiently positive to turn to formal econometric analysis. The non-stationary character of many of the series requires us to first establish unit root type non-stationarity in order to be able to use (panel) cointegration analysis to test for 'long-run' relationships. We turn to both of these issues in the subsequent two sections.

4 Econometric Analysis I: Panel Unit Root Testing

The small sample size with only nine years necessitates the application of panel unit root tests. The implemented panel unit root tests are developed in the following papers:[11] Levin,

[11] As indicated already above, the implementation of the econometric procedures was originally based on Chiang and Kao (2002), but has been substantially modified, *corrected* and extended. The authors currently

Lin and Chu (2002), abbreviated by LL; Breitung (2000), abbreviated by UB; two tests developed in Im, Pesaran and Shin (1997) and Im, Pesaran and Shin (2003), a t-type test abbreviated by IPS and a Lagrange multiplier test, abbreviated by $IPS - LM$; Harris and Tzavalis (1999), abbreviated by HT; and Maddala and Wu (1999), abbreviated by MW.

All tests except for HT allow for individual specific serial correlation structures, whilst HT is designed for the case of no serial correlation in the residuals. For all tests the null hypothesis is that of a unit root in all series. The alternative is stationarity in all series, except for the tests developed by Im et al. where the alternative allows for non-stationarity in a non-vanishing (in the limit for $N \to \infty$) fraction of the series. The first five tests listed above are asymptotically normally distributed and the latter is asymptotically χ^2_{2N} distributed, with N denoting the cross section dimension of the panel. The test LL, UB, IPS and HT are left-sided and $IPS - LM$ and MW are right sided.

It has been found, see e.g. Hlouskova and Wagner (2004a,b), that for panels of the size available in this study, the asymptotic distributions of the panel unit root and panel cointegration tests provide poor approximations to the small sample distributions (for an example see Figure 3 and the corresponding discussion below). In other words, the notorious size and power problems for which unit root tests are known in the time series context also appear in small or short panels.

To overcome these limitations we have implemented three different bootstrap methods to obtain improved small sample inference. The three bootstrap methods, explained in Appendix C, are the *parametric*, the *non-parametric* and the *residual based block* bootstrap. The latter has been developed for non-stationary time series by Paparoditis and Politis (2003). The former two methods obtain white noise bootstrap replications of residuals due to pre-whitening and the latter is based on re-sampling blocks of residuals to preserve the serial correlation structure. The difference between the parametric and the non-parametric bootstrap is essentially given by the fact that in the former the residuals are drawn from a normal distribution and are re-sampled from the empirical residuals in the latter. The results obtained by the three bootstrap methods are rather similar, thus in the main text we only report the result from one of the methods. Note furthermore that bootstrapping, if re-sampling is done identically for all cross-sectional units, also provides certain robustness against the vio-

work on a user friendly version of the new toolbox. A detailed description of these panel unit root tests, including the assumptions on the data generating process and the precise construction of the test statistics, is given e.g. in Hlouskova and Wagner (2004a).

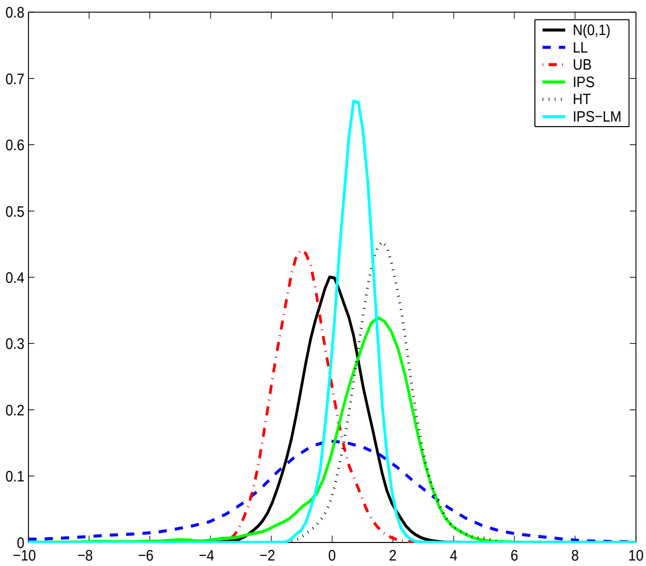

Figure 3: Bootstrap test statistic distributions for relative prices, p^{rel}, for the five asymptotically standard normally distributed panel unit root tests.
The results are based on the non-parametric bootstrap with 5000 replications. Fixed effects are included.

lation of a key assumption of all the implemented unit root tests, namely the assumption of cross-sectional independence, see e.g. Chang (2000). In Figure 3 we display the asymptotic null distribution (the standard normal distribution) and the bootstrap null distributions (from the non-parametric bootstrap) for one of the variables tested for a unit root, the relative price of non-tradables to tradables, p^{rel}, for the five asymptotically standard normally distributed tests. The figure shows *substantial* differences between the bootstrap approximations to the finite sample distribution of the tests and their asymptotic distribution. Thus, basing inference on the asymptotic critical values leads to substantial size distortions. This can also be seen in Tables 2 and 3 below, where in brackets the bootstrap 5% critical values are displayed. They vary substantially both across tests and also across variables, and are in many cases far away from the asymptotic critical values ± 1.645, respectively 26.296 for the Maddala and Wu test.

All tests are implemented, as is standard in the unit root literature, with different specifications of the deterministic terms. We have tested with two versions, for each of the three bootstrap algorithms. With fixed effects only, reported in Table 2 and with fixed effects and individual specific linear trends, reported in Table 3. These tables report the results based on the parametric bootstrap. All bootstrap results in this paper are based on 5000 replications. Including a linear trend in the test equation, when there is no trend in the data generating process reduces the power of the tests, on the other hand, omitting a linear trend when there is a trend in the data, induces a bias in the tests towards the null hypothesis. Graphical inspection of the data shows that for basically all variables, even the relative price or wage variables, in at least one or two countries trending behavior is visible. This implies that in the panel framework the specification with trends may be more appropriate. As is common in the unit root literature, we however present and compare the results of both specifications for all variables. The comparison of the results obtained with different specifications is usually informative.

The variables for which we report the panel unit root test results are the following.[12] Three real exchange rate measures, q the logarithm of the real exchange of the CEEC countries to the EU11 (indexed to equal 100 in 1995) based on the GDP deflators; q_2 defined similarly to q but based on the price indices of only the tradables and non-tradables sectors; and q^T, the real exchange rate for tradables (again in logarithms and indexed). The latter is directly related to one of the assumptions of the standard Baumol-Bowen and Balassa-Samuelson models, namely prevalence of PPP in the tradables sector. The econometric testing for validity of PPP in a world of I(1) nonstationary data is to test for stationarity of the real exchange rate. This, of course, allows for substantial and persistent differences in prices. The unit root hypothesis is hardly at all rejected for these variables, in particular if one relies on bootstrap based inference, then only for q_2 two tests reject the null when a trend is included. In particular also note that for q^T some rejections occur based on the asymptotic critical values, but no rejection occurs based on the bootstrap critical values. Thus, we conclude that PPP does not hold in the tradables sector between the CEE countries and the EU11.

The second group of variables tested are the various (logarithms of) price variables. The relative price of non-tradables to tradables and different price level differentials between the CEE countries and the EU11. For the price level differentials it is not a priori clear which specification is preferable, since due to catching-up of the CEE countries persistent inflation differentials and thus a narrowing of price differentials might induce a trend in the price level

[12]See also Table 23 for a summary description of the variables and transformations. Note that also for the output variables and the prices the null hypothesis of a unit root can generally not be rejected, detailed results are available from the authors upon request. In the presentation here we focus on those variables and their relationships that are directly relevant for the model only.

differences.

Concerning the relative price, p^{rel}, the hypothesis of a unit root is never rejected in both specifications. For the price level differentials the evidence is a bit more mixed. The tests IPS and MW reject, based on the bootstrap critical values, the null of a unit root for all three price level differentials, $p^T - p^{T*}$, $p^{GDP} - p^{GDP*}$ and $p^{T+N} - p^{(T+N)*}$. Also $IPS - LM$ is rejecting the null for these three variables. When a linear trend is included in the test equation, for $p^{GDP} - p^{GDP*}$ three tests reject the null. Thus, for the price differentials some evidence for stationarity is available.

The third group of variables are four wage variables, again normalized to 100 in 1995 and in logarithms. We have tested the wages in the tradables sector, w^T, the wages in the non-tradables sector, w^N and the relative wage in the non-tradables to the tradables sector, w^{rel}. Additionally we also test for a unit root in the variable $w^{rel}_{BS} = w^{rel} - w^{rel*} + \ln 100$.[13] This latter variable plays, up to neglected constants $\delta\alpha^N$ a role in the extended Balassa-Samuelson model, compare equation (10). For the sectoral wages, the specification with trend in the test equation seems to be more relevant. As expected, none of the tests rejects the null of a unit root in these two variables. Given the unit root non-stationarity of w^N and w^T testing for a unit root in w^{rel} is obviously a direct device of testing for cointegration of the form $[1, -1]$ between the wages in the two sectors. Thus, similarly to PPP above, stationarity of relative wages is a weak econometric formulation of *wage homogeneity*. A unit root in w^{rel} is not rejected, when the bootstrap critical values are employed, with one exception, see again Tables 2 and 3. w^{rel} is one of the examples where inference based on the asymptotic critical values leads, for some tests at least, to the incorrect conclusion of rejecting the null of a unit root. Thus, we conclude that the assumption of wage homogeneity is not fulfilled in the CEECs. Also for the variable w^{rel}_{BS} the rejections of the unit root hypothesis stem from applying asymptotic critical values. With bootstrap critical values only the HT test with no trends rejects the null hypothesis. Thus, also w^{rel}_{BS} is found to be non-stationary.

Next, the productivity variables are tested. Again, there a four variables considered, normalized to 100 in 1995 and in logarithms: Productivity in the tradables sector, a^T; in the non-tradables sector, a^N; relative productivity in the tradables to the non-tradables sector a^{rel} and the differential of relative productivities in the CEE country and the EU11, $a^{rel} - a^{rel*}$. The latter is, as discussed above, a widely used variable in the BS models, see equations (8)

[13] The factor $\ln 100$ is added to achieve that the variable w^{rel}_{BS} equals $\ln 100$ in 1995.

or (10). The results are as follows. Only for a^T and the inappropriate specification without trends, several tests reject the null of a unit root. For the other three measures only one or two tests reject. The difference in relative productivities is thus, with two rejections in the specification with trends, sort of a borderline case.

Finally the other BS variables discussed in Section 2, $BSE1$, $BSE2$ and BS, are tested for a unit root. The evidence for all these variables is rather clear. The unit root hypothesis is never rejected for $BSE1$ and $BSE2$. For BS two rejections occur in the specification with trend. Twice a unit root is also rejected for BS when using the asymptotic critical values. Thus, only for the relative productive differentials weighted by the δ's there is at least some evidence for stationarity.

The unit root testing performed in this section leads to two main conclusions. First, unit root non-stationarity prevails throughout the variables. Second, no evidence for PPP in tradables between the CEECs and the EU11 is found. Also relative wages are found to be non-stationary in the CEECs. These two facts imply that the empirical analysis has to focus on specifications that do not rely upon PPP in the tradables sector and that do not rely upon homogenous wages, i.e. the so called extended models form the basis for subsequent analysis. The next step, given the unit root non-stationary behavior is to test for cointegration. This is done in the following section.

	LL	UB	IPS	HT	$IPS-LM$	MW
q	**-2.593*** (-5.626)	0.906 (-1.560)	**-2.348** (-2.190)	-0.048 (-0.150)	1.048 (1.345)	**62.572*** (64.125)
q_2	**-2.462*** (-5.603)	1.131 (-1.692)	-0.905 (-2.362)	0.244 (-0.258)	0.272 (1.403)	**36.435*** (65.853)
q^T	-1.083 (-5.777)	0.294 (-1.630)	**-1.660*** (-2.193)	-0.254 (-1.298)	0.916 (1.364)	**39.250*** (62.507)
p^{rel}	3.235 (-5.231)	1.538 (-0.454)	0.212 (-1.172)	0.021 (-0.005)	-0.465 (0.621)	19.091 (54.766)
$p^T - p^{T*}$	**-4.001*** (-5.639)	0.509 (-1.724)	**-4.902** (-2.311)	0.303 (0.388)	**2.138** (1.405)	**121.282** (68.353)
$p^{GDP} - p^{GDP*}$	1.474 (-5.535)	-0.983 (-1.726)	**-5.177** (-2.287)	**0.161** (0.831)	**2.382** (1.435)	**121.451** (69.071)
$p^{T+N} - p^{(T+N)*}$	**-2.606*** (-5.572)	-0.151 (-1.722)	**-4.334** (-2.221)	**0.475** (0.964)	**2.189** (1.415)	**94.208** (68.859)
w^T	**-3.371*** (-5.792)	1.521 (-1.699)	**-2.022*** (-2.137)	0.891 (0.794)	**1.481** (1.420)	**39.818*** (68.135)
w^N	**-9.505** (-4.744)	1.927 (-1.650)	**-1.652*** (-2.024)	1.386 (0.882)	1.039 (1.369)	**52.027*** (68.451)
w^{rel}	-1.109 (-5.074)	-0.807 (-1.369)	-0.937 (-2.085)	**-4.210*** (-4.847)	0.484 (1.258)	**33.806*** (54.748)
w^{rel}_{BS}	**-1.656*** (-4.220)	**-1.133** (-0.823)	-0.726 (-2.876)	**-4.236** (-3.532)	0.353 (0.966)	**32.950*** (68.980)
a^T	-0.925 (-5.388)	0.820 (-1.770)	1.471 (-2.380)	1.823 (0.384)	-1.724 (1.421)	7.510 (67.647)
a^N	**-2.066*** (-5.358)	1.573 (-1.609)	3.362 (-2.188)	2.407 (-0.512)	-2.373 (1.373)	2.756 (63.215)
a^{rel}	**-2.658*** (-5.323)	0.084 (-1.666)	-0.488 (-2.292)	0.039 (-1.438)	0.264 (1.334)	18.226 (60.764)
$a^{rel} - a^{rel*}$	**-2.658*** (-5.323)	-0.287 (-1.631)	**-2.898** (-2.238)	-0.529 (-2.067)	1.247 (1.330)	**77.414** (60.045)
BS	4.468 (-5.155)	-1.227 (-1.600)	1.154 (-2.263)	-1.619 (-3.398)	-1.564 (1.331)	6.869 (60.986)
$BSE1$	1.210 (-5.270)	0.579 (-1.619)	1.978 (-2.202)	0.745 (-2.057)	-2.025 (1.329)	5.527 (59.346)
$BSE2$	1.106 (-5.465)	0.492 (-1.638)	2.906 (-2.211)	0.363 (-1.806)	-1.793 (1.336)	5.581 (59.394)

Table 2: Results of panel unit root tests including fixed effects. In parentheses the 5% critical values obtained by the *parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by -1.645 for the first 4 tests, by 1.645 for IPS-LM and by 26.296 for MW.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both the autoregression based tests and the parametric bootstrap are equal to 1.

	LL	UB	IPS	HT	$IPS-LM$	MW
q	0.650 (-10.323)	-0.544 (-0.957)	-1.111 (-3.148)	1.122 (-1.509)	0.180 (0.802)	**41.843*** (94.445)
q_2	1.945 (-10.053)	-0.436 (-0.931)	**-3.249** (-2.934)	0.752 (-2.202)	0.420 (0.849)	**109.131** (89.893)
q^T	-0.862 (-10.143)	-0.852 (-0.930)	**-1.819*** (-2.956)	0.088 (-2.365)	0.684 (0.861)	**47.933*** (90.326)
p^{rel}	0.334 (-10.003)	-0.664 (-0.906)	1.328 (-3.627)	-1.455 (-4.393)	-1.834 (0.844)	13.641 (99.438)
$p^T - p^{T*}$	**-2.481*** (-10.847)	0.114 (-0.895)	-1.035 (-3.029)	2.429 (-0.495)	0.488 (0.848)	**30.208*** (97.007)
$p^{GDP} - p^{GDP*}$	5.423 (-10.317)	**-1.671** (-0.969)	**-7.038** (-3.468)	2.871 (0.287)	**0.890** (0.826)	**53.705*** (103.498)
$p^{T+N} - p^{(T+N)*}$	0.784 (-10.536)	0.071 (-0.934)	**-4.085** (-3.294)	2.993 (0.178)	0.746 (0.826)	**119.656** (99.954)
w^T	0.812 (-11.082)	0.068 (-0.963)	0.411 (-3.273)	2.453 (-0.961)	-0.954 (0.855)	20.508 (97.779)
w^N	-1.192 (-9.684)	0.864 (-0.959)	0.769 (-3.186)	2.600 (-1.662)	-1.033 (0.877)	13.311 (97.199)
w^{rel}	**-6.616*** (-9.427)	0.627 (-0.920)	**-4.195** (-2.779)	**-2.462*** (-4.286)	-0.047 (0.791)	**48.768*** (86.720)
w^{rel}_{BS}	**-5.343*** (-8.495)	0.473 (-0.928)	**-2.267*** (-5.015)	**-2.507*** (-4.424)	-0.328 (0.815)	**94.633*** (107.882)
a^T	**-13.737** (-10.469)	-0.661 (-0.908)	**-4.166** (-2.858)	-0.425 (-2.998)	**1.083** (0.849)	**117.895** (88.316)
a^N	0.372 (-10.054)	-0.192 (-0.916)	1.030 (-2.812)	-0.807 (-3.712)	-1.187 (0.804)	9.801 (86.635)
a^{rel}	**-11.985** (-10.589)	-0.829 (-0.917)	-1.280 (-2.878)	0.220 (-2.984)	-0.009 (0.838)	**52.175*** (88.940)
$a^{rel} - a^{rel*}$	**-11.985** (-10.589)	0.299 (-0.832)	**-2.043*** (-2.854)	0.152 (-2.849)	0.355 (0.850)	**68.918*** (89.618)
BS	**-8.828*** (-8.913)	0.357 (-0.925)	**-5.126** (-2.955)	**-2.081*** (-4.504)	0.617 (0.790)	**157.148** (90.981)
$BSE1$	3.468 (-9.636)	-0.180 (-0.903)	0.291 (-2.867)	-0.194 (-3.043)	-1.089 (0.845)	**30.413*** (90.354)
$BSE2$	-0.120 (-10.100)	0.472 (-0.872)	1.158 (-3.029)	0.073 (-2.525)	-0.934 (0.856)	8.957 (94.265)

Table 3: Results of panel unit root tests including fixed effects and time trends. In parentheses the 5% critical values obtained by the *parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by -1.645 for the first 4 tests, by 1.645 for IPS-LM and by 26.296 for MW.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both the autoregression based tests and the parametric bootstrap are equal to 1.

5 Econometric Analysis II: Panel Cointegration Testing

In total ten cointegration tests are performed, seven of them developed in Pedroni (2004) and three developed in Kao (1999). Similar bootstrap procedures as for the panel unit root tests are applied, for details see again Appendix C.

All employed tests have the null hypothesis of no cointegration and are based on the residuals of the so called *cointegrating* regression. Thus, the null hypothesis of no cointegration is equivalent to a unit root in the residuals of the cointegrating regression. The usual specifications concerning deterministic variables have been implemented. We report again test results when including only fixed effects, and when including fixed effects and individual specific trends.

Pedroni (2004) develops four *pooled* tests and three *group-mean* tests. Three of the four pooled tests are based on a first order autoregression and correction factors in the spirit of Phillips and Ouliaris (1990). These are a variance-ratio statistic, PP_σ; a test statistic based on the estimated first-order correlation coefficient, PP_ρ; and a test based on the t-value of the first-order correlation coefficient, PP_t. The fourth test is based on an augmented Dickey-Fuller type test statistic, PP_{df}, in which the correction for serial correlation is achieved by augmenting the test equation by lagged differenced residuals of the cointegrating regression. Thus, this test is a panel cointegration analogue of the panel unit root test of Levin, Lin and Chu (2002) discussed above. For these four tests the alternative hypothesis is stationarity with a homogeneity restriction on the first order correlation in all cross-section units.

To allow for a slightly less restrictive alternative Pedroni (2004) develops three group-mean tests. For these tests the alternative allows for completely heterogeneous correlation patterns in the different cross-section units. The group-mean tests can be seen as averaged - over the cross-section units - test statistics. Pedroni discusses the group-mean analogues of all but the variance-ratio test statistic. Paralleling the above notation for the pooled tests, we denote them with PG_ρ, PG_t and PG_{df}.

After centering and scaling the test statistics by suitable correction factors, to correct for serial correlation of the residuals and for potential endogeneity of the regressors in the cointegrating regression, all test statistics are asymptotically standard normally distributed. The first test, PP_σ, is right-sided and the other six tests are left-sided.

Kao (1999) derives tests similar to three of the pooled tests of Pedroni for *homogenous* panels and when only fixed effects are included. A panel is called homogenous, if the serial correlation pattern is identical across units. Kao's three tests, K_ρ, K_t and K_{df} using obvious abbreviations, are based on the spurious least squares dummy variable (LSDV) estimator of the cointegrating regression. We report results obtained by these tests, in Appendix B. We include these tests, because it might be the case that in small samples tests based on a cross-sectional homogeneity assumption perform comparatively well, since no individual specific correlation corrections, which may be very inaccurate in short panels, are necessary. Also Kao's tests are after scaling and centering appropriately asymptotically standard normally distributed and left sided.

Figures similar to Figure 3 are available from the authors upon request. Again substantial differences between the asymptotic critical values and the bootstrap critical values emerge. Note also again that bootstrapping robustifies, when done identically for all cross-section units, the tests to a certain extent against a violation of the critical assumption of cross-sectional independence, which is required for all tests discussed. We report in Tables 5 and 6 the results obtained by applying the non-parametric bootstrap, for a list of relationships discussed next and summarized in Table 4.

Label	Relationships tested for cointegration	FM-OLS	D-OLS
Wages	$w_{it}^{T} = c_i + \beta w_{it}^{N} + u_{it}$	+	+
LC-LPT	$lcr_{it}^{T} = c_i + \beta a_{it}^{T} + u_{it}$	+	+
LC-LPN	$lcr_{it}^{N} = c_i + \beta a_{it}^{N} + u_{it}$	+	+
BBE	$p_{it}^{rel} = c_i + \beta_1 a_{it}^{rel} + \beta_2 w_{it}^{rel} + u_{it}$	+	+
Aq	$q_{it} = c_i + \beta_1 q_{it}^{T} + \beta_2(\delta_{it}(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - \delta_t^*(a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	−
Aq_2	$q_{2,it} = c_i + \beta_1 q_{it}^{T} + \beta_2(\delta_{it}(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - \delta_t^*(a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	+
Ap	$p_{it}^{GDP} - p_t^{GDP*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(\delta_{it}(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - \delta_t^*(a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	+
Ap_2	$p_{it}^{T+N} - p_t^{(T+N)*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(\delta_{it}(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - \delta_t^*(a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	+
Bq	$q_{it} = c_i + \beta_1 q_{it}^{T} + \beta_2(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel} - (a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	+
Bq_2	$q_{2,it} = c_i + \beta_1 q_{it}^{T} + \beta_2(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel} - (a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	+
Bp	$p_{it}^{GDP} - p_t^{GDP*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2((a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel} - (a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	+	+
Bp_2	$p_{it}^{T+N} - p_t^{(T+N)*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(a_{it}^{rel} + \alpha_{it}^{N} w_{it}^{rel}) - (a_t^{rel*} + \alpha_t^{N*} w_t^{rel*}) + u_{it}$	+	+
Cq	$q_{it} = c_i + \beta_1 q_{it}^{T} + \beta_2(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	−
Cq_2	$q_{2,it} = c_i + \beta_1 q_{it}^{T} + \beta_2(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	+
Cp	$p_{it}^{GDP} - p_t^{GDP*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	+
Cp_2	$p_{it}^{T+N} - p_t^{(T+N)*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	+
Dq	$q_{it} = c_i + \beta_1 q_{it}^{T} + \beta_2(a_{it}^{rel} - a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	−	−
Dq_2	$q_{2,it} = c_i + \beta_1 q_{it}^{T} + \beta_2(a_{it}^{rel} - a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	+
Dp	$p_{it}^{GDP} - p_t^{GDP*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(a_{it}^{rel} - a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	−
Dp_2	$p_{it}^{T+N} - p_t^{(T+N)*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2(a_{it}^{rel} - a_t^{rel*}) + \beta_3 w_{BS,it}^{rel} + u_{it}$	+	+
Eq	$q_{it} = c_i + \beta_1 q_{it}^{T} + \beta_2 a_{it}^{rel} + \beta_3 a_t^{rel*} + \beta_4 w_{BS,it}^{rel} + u_{it}$	−	−
Eq_2	$q_{2,it} = c_i + \beta_1 q_{it}^{T} + \beta_2 a_{it}^{rel} + \beta_3 a_t^{rel*} + \beta_4 w_{BS,it}^{rel} + u_{it}$	−	−
Ep	$p_{it}^{GDP} - p_t^{GDP*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2 a_{it}^{rel} + \beta_3 a_t^{rel*} + \beta_4 w_{BS,it}^{rel} + u_{it}$	−	−
Ep_2	$p_{it}^{T+N} - p_t^{(T+N)*} = c_i + \beta_1(p_{it}^{T} - p_t^{T*}) + \beta_2 a_{it}^{rel} + \beta_3 a_t^{rel*} + \beta_4 w_{BS,it}^{rel} + u_{it}$	−	+

Table 4: Relationships tested for cointegration. In the table c_i equals β_{0i} if only fixed effects are included and $c_i = \beta_{0i} + \beta_i t$ if fixed effects and time trends are included in the cointegrating regression.

Columns FM-OLS and D-OLS display the information concerning the signs of the estimated coefficients (including only fixed effects) obtained by applying FM-OLS and D-OLS respectively. Here '+' indicates that all coefficient signs are in accordance with the theoretical model, '-' indicates that at least one sign is not according to the model.

The first relationship investigated is again the relationship between sectoral wages. In the previous section non-stationarity of w^{rel} could not be rejected. In this section we search for less restricted cointegration, i.e. we test for cointegration in $w_{it}^T = c_i + \beta w_{it}^N + u_{it}$. Throughout the section c_i denotes either fixed effects or fixed effects and individual specific time trends, depending upon the specification investigated. The result is interesting, as all Pedroni tests reject the null of no cointegration (based on bootstrap critical values) in case a linear trend is included in the unit root test regression. Thus, relative wages are found to be cointegrated when allowing for a linear trend in the cointegrating relationship and when the coefficient β is not restricted to equal 1. This, however, is a relatively weak link in wages across the two sectors, which we certainly do not interpret as evidence for wage homogeneity.

The second and third relationship, LC-LPT and LC-LPN, are included to verify one of the underlying assumptions of the Baumol-Bowen and Balassa-Samuelson model: competitive wage setting and the implied link between productivity and total labor costs in the tradables and the non-tradables sector respectively. For a Cobb-Douglas production function the marginal product equals the average product. Thus, if wages are set competitively, wages equal the marginal and in the Cobb-Douglas case thus also the average product of labor. Therefore, a weak empirical formulation of this relationship is cointegration between (average) labor productivity and total labor costs. The evidence for cointegration is rather weak for both sectors. However, compared to other relationships, for these relationships at least some tests reject the null of no cointegration. Thus, a link between (log levels of) labor productivities and labor costs is not entirely rejected at least.

	PP_σ	PP_ρ	PP_t	PP_{df}	PG_ρ	PG_t	PG_{df}
$Wages$	**3.228** (2.728)	-1.532 (-1.905)	**-3.169*** (-4.413)	**-2.619*** (-3.979)	-0.160 (-0.297)	**-3.224*** (-4.307)	**-3.666*** (-5.241)
LC-LPT	1.053 (1.589)	-1.143 (-2.165)	**-2.666*** (-5.320)	**-2.258*** (-4.738)	0.196 (-0.591)	**-2.813*** (-6.030)	**-2.977*** (-6.467)
LC-LPN	1.549 (2.433)	-1.399 (-1.865)	**-4.463** (-4.430)	**-3.758*** (-4.043)	0.091 (-0.291)	**-5.107** (-4.563)	**-5.317** (-5.290)
BBE	-0.414 (0.286)	0.369 (-0.384)	-0.836 (-4.179)	-0.430 (-3.638)	1.403 (0.957)	-1.266 (-4.773)	**-1.677*** (-5.049)
Aq	-1.324 (-0.338)	1.271 (-0.384)	-0.008 (-4.277)	0.516 (-3.895)	1.718 (0.940)	-0.981 (-4.791)	-1.524 (-5.367)
Aq_2	-1.610 (-0.253)	1.448 (-0.365)	0.489 (-4.239)	0.963 (-3.789)	1.855 (0.999)	-0.764 (-4.427)	-1.274 (-4.966)
Ap	-1.113 (0.532)	0.752 (-0.322)	-0.899 (-4.037)	-0.579 (-3.680)	1.397 (1.038)	**-1.778*** (-4.329)	**-2.406*** (-5.026)
Ap_2	-1.238 (0.770)	0.775 (-0.276)	-0.714 (-3.869)	-0.442 (-3.459)	1.310 (1.026)	**-1.957*** (-4.501)	**-2.588*** (-5.074)
Bq	-1.878 (-0.609)	1.120 (-0.438)	-0.381 (-4.364)	0.296 (-3.995)	1.872 (0.988)	-0.377 (-4.549)	-0.828 (-5.167)
Bq_2	-2.098 (-0.539)	1.616 (-0.429)	0.775 (-4.313)	1.370 (-3.903)	1.969 (0.972)	0.683 (-4.612)	-0.322 (-5.292)
Bp	-0.097 (0.660)	-0.340 (-0.520)	**-3.559*** (-4.568)	**-3.233*** (-4.167)	**0.814** (0.906)	**-3.376*** (-4.692)	**-4.096*** (-5.510)
Bp_2	-0.211 (0.893)	-0.297 (-0.567)	**-2.957*** (-4.821)	**-2.815*** (-4.242)	**0.773** (0.901)	**-2.950*** (-4.883)	**-3.704*** (-5.639)
Cq	-2.840 (0.169)	2.061 (0.419)	-0.682 (-5.061)	-0.382 (-4.360)	2.523 (1.870)	-0.909 (-5.358)	-1.451 (-5.801)
Cq_2	-2.839 (0.110)	2.032 (0.414)	-0.594 (-5.194)	-0.405 (-4.394)	2.680 (1.865)	-0.342 (-5.276)	-0.868 (-5.763)
Cp	-0.666 (0.781)	0.646 (0.348)	**-3.224*** (-5.487)	**-2.989*** (-4.736)	1.908 (1.714)	**-3.673*** (-6.484)	**-4.633*** (-6.956)
Cp_2	-0.747 (0.832)	0.682 (0.327)	**-3.124*** (-5.481)	**-2.952*** (-4.791)	1.910 (1.708)	**-3.576*** (-6.337)	**-4.458*** (-6.844)
Dq	-1.418 (-0.362)	1.627 (0.637)	-0.754 (-4.129)	0.406 (-3.621)	2.336 (2.065)	-1.063 (-4.234)	**-1.723*** (-4.724)
Dq_2	-1.329 (-0.162)	1.752 (0.669)	-0.250 (-3.999)	0.790 (-3.432)	2.354 (2.090)	-1.252 (-4.117)	**-1.891*** (-4.683)
Dp	-1.186 (0.485)	1.711 (0.535)	-0.324 (-4.539)	0.826 (-4.049)	2.172 (1.902)	**-2.260*** (-5.109)	**-3.042*** (-5.678)
Dp_2	-1.589 (0.492)	1.899 (0.540)	0.163 (-4.541)	1.400 (-3.990)	2.134 (1.918)	**-2.786*** (-4.827)	**-3.480*** (-5.474)
Eq	-0.354 (0.393)	1.759 (1.531)	**-1.811*** (-3.861)	-1.265 (-3.436)	3.038 (2.853)	**-2.227*** (-4.334)	**-2.430*** (-5.329)
Eq_2	-0.534 (0.394)	1.574 (1.495)	**-2.537*** (-4.065)	**-2.220*** (-3.702)	2.894 (2.826)	**-2.453*** (-4.694)	**-2.957*** (-5.514)
Ep	0.010 (0.380)	**1.347** (1.363)	**-4.172*** (-5.166)	**-3.782*** (-4.511)	**2.639** (2.713)	**-4.256*** (-5.860)	**-4.740*** (-6.420)
Ep_2	-0.723 (0.374)	**1.207** (1.406)	**-4.660*** (-4.817)	**-4.646** (-4.188)	**2.574** (2.760)	**-4.842*** (-5.477)	**-5.977*** (-6.166)

Table 5: Results of Pedroni's panel cointegration tests including fixed effects. In parentheses the 5% critical values obtained by the *non-parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by 1.645 for the first test and by -1.645 for the other 6 tests.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both the autoregression based tests and the non-parametric bootstrap are equal to one. The window-length of the Bartlett kernel for the non-parametric tests is also equal to one.

	PP_σ	PP_ρ	PP_t	PP_{df}	PG_ρ	PG_t	PG_{df}
$Wages$	**0.922** (0.749)	**-0.551** (-0.170)	**-6.123** (-4.898)	**-5.681** (-4.474)	**0.762** (1.077)	**-6.317** (-4.824)	**-6.628** (-5.784)
LC-LPT	-0.313 (0.650)	0.636 (-0.158)	**-2.287*** (-4.865)	**-1.768*** (-4.367)	1.711 (1.124)	**-1.927*** (-4.648)	**-2.482*** (-5.631)
LC-LPN	-0.405 (0.611)	0.523 (-0.064)	**-2.761*** (-4.558)	**-2.452*** (-4.081)	1.727 (1.191)	**-2.456*** (-4.285)	**-4.170***(-5.336)
BBE	-0.540 (0.023)	1.120 (0.919)	**-3.401*** (-4.467)	**-3.120*** (-3.990)	2.258 (2.082)	**-3.386*** (-4.442)	**-3.858*** (-5.370)
Aq	-0.480 (-0.020)	1.165 (0.888)	**-2.794*** (-4.615)	**-2.264*** (-4.130)	2.276 (2.051)	**-2.365*** (-4.802)	**-3.045*** (-5.742)
Aq_2	-0.461 (0.005)	1.399 (0.881)	-1.547 (-4.602)	-1.075 (-4.072)	2.504 (2.050)	-1.187 (-4.811)	**-1.759*** (-5.708)
Ap	-0.536 (0.052)	1.321 (0.848)	**-2.773*** (-4.847)	**-1.983*** (-4.301)	2.359 (2.007)	**-2.857*** (-5.057)	**-3.382*** (-5.892)
Ap_2	-0.476 (0.054)	1.314 (0.810)	**-2.037*** (-5.098)	**-1.768*** (-4.443)	2.415 (1.978)	-1.349 (-5.217)	**-2.135*** (-5.984)
Bq	-0.520 (0.030)	1.341 (0.858)	**-2.187*** (-4.703)	**-1.793*** (-4.178)	2.467 (2.010)	-1.580 (-4.822)	**-2.133*** (-5.441)
Bq_2	-0.464 (0.057)	1.366 (0.849)	-1.624 (-4.798)	-1.462 (-4.311)	2.521 (2.013)	-0.540 (-4.863)	-1.196 (-5.639)
Bp	-0.067 (0.076)	0.980 (0.820)	**-3.674*** (-4.910)	**-3.371*** (-4.487)	2.144 (1.977)	**-2.899*** (-5.084)	**-3.571*** (-5.929)
Bp_2	-0.181 (0.099)	**0.739** (0.823)	**-4.226*** (-4.929)	**-4.124*** (-4.468)	**1.921** (1.981)	**-3.002*** (-5.077)	**-3.892*** (-6.100)
Cq	-1.441 (-0.585)	2.183 (1.736)	**-1.977*** (-4.651)	**-1.891*** (-4.145)	3.169 (2.883)	-1.318 (-5.019)	**-2.257*** (-5.971)
Cq_2	-1.508 (-0.568)	2.179 (1.739)	**-2.287*** (-4.693)	-1.318 (-4.693)	3.162 (2.885)	**-2.031*** (-4.996)	**-2.699*** (-5.875)
Cp	-1.201 (-0.480)	2.391 (1.713)	-0.730 (-4.903)	-0.272 (-4.355)	3.411 (2.847)	0.206 (-5.297)	-0.527 (-6.091)
Cp_2	-1.203 (-0.476)	2.385 (1.682)	-0.804 (-5.070)	-0.566 (-4.458)	3.458 (2.822)	0.292 (-5.289)	-0.449 (-6.134)
Dq	-1.123 (-0.537)	2.003 (1.771)	**-2.507*** (-4.304)	**-1.963*** (-3.804)	3.150 (2.908)	-1.583 (-4.407)	**-2.322*** (-5.424)
Dq_2	-1.059 (-0.545)	2.126 (1.790)	-1.558 (-4.216)	-1.429 (-3.785)	3.262 (2.924)	-0.363 (-4.384)	-1.119 (-5.324)
Dp	-0.755 (-0.443)	1.989 (1.632)	**-2.656*** (-5.344)	**-2.272*** (-4.680)	3.119 (2.777)	**-1.847*** (-5.630)	**-2.622*** (-6.237)
Dp_2	-0.787 (-0.433)	1.926 (1.661)	**-2.899*** (-5.108)	**-2.657*** (-4.480)	3.035 (2.794)	**-2.170*** (-5.457)	**-2.957*** (-6.144)
Eq	-1.995 (-1.002)	2.930 (2.471)	-1.005 (-4.503)	-0.346 (-3.807)	4.046 (3.614)	-0.464 (-4.935)	-1.232 (-5.942)
Eq_2	-2.005 (-1.009)	2.769 (2.464)	**-1.825*** (-4.559)	-1.125 (-3.949)	3.893 (3.612)	**-1.795*** (-5.022)	**-2.762*** (-6.003)
Ep	-1.088 (-0.950)	2.685 (2.425)	**-2.548*** (-4.977)	**-1.940*** (-4.296)	3.805 (3.573)	**-2.257*** (-5.395)	**-3.034*** (-6.392)
Ep_2	-1.205 (-0.971)	**2.401** (2.434)	**-4.763*** (-4.918)	**-3.798*** (-4.211)	**3.496** (3.576)	**-5.384*** (-5.434)	**-6.420** (-6.266)

Table 6: Results of Pedroni's panel cointegration tests including fixed effects and time trends. In parentheses the 5% critical values obtained by the *non-parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by 1.645 for the first test and by -1.645 for the other 6 tests.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both the autoregression based tests and the non-parametric bootstrap are equal to one. The window-length of the Bartlett kernel for the non-parametric tests is also equal to one.

The fourth relationship, BBE, tests for cointegration in the extended Baumol-Bowen model (4) (with the discussed focus on $a^T - a^N$ instead of $\frac{\alpha^N}{\alpha^N} a^T - a^N$), i.e. testing is performed on the equation $p_{it}^{rel} = c_i + \beta_1 a_{it}^{rel} + \beta_2 w_{it}^{rel}$. Based on bootstrap inference, no evidence for cointegration is found. This relationship is again an example where inference based on the asymptotic critical values leads to the incorrect conclusion of cointegration, in particular in the specification including trends.

The remaining twenty relationships tested are structured as follows. The first letter in the name, ranging from A to E, indicates the specification of the *Balassa-Samuelson* term or variable. The remaining one respectively two symbols in the name indicate the dependent variable, q and q_2, the two real exchange rate measures, and p and p_2, when the dependent variables are the respective price differentials between the CEE country and the EU11. Equations A include BSE1 (defined in Section 2) as their BS variable, equations B include BSE2 as their BS variable. In the equations labelled C and D, the wage components of the BS variable are treated separately. Thus, in the C equations, the BS variable is given by $\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}$, and in the D equations by $a_{it}^{rel} - a_t^{rel*}$. In both sets of equations, C and D, wages are included in the form of w_{BS}^{rel}, introduced already in the previous section. Finally the E equations relax the homogeneity assumption on the productivity terms. They nest equations D. The corresponding parameter restrictions can be tested. The Balassa-Samuelson variable in the E equations is a combination of the relative productivities at home and in the EU11. As discussed above, four dependent variables are chosen for the equations, two price differentials, $p_{it}^{GDP} - p_t^{GDP*}$ and $p_{it}^{T+N} - p_t^{(T+N)*}$ and two real exchange rate variables, $q_{it} = e_{it} + p_t^{GDP*} - p_{it}^{GDP}$ and $q_{2,it} = e_{it} + p_t^{(T+N)*} - p_{it}^{T+N}$.

The test results, in Tables 5 and 6 for the discussed twenty Balassa-Samuelson relationships are very clear. There is basically no evidence for cointegration, if one bases inference on any of the implemented bootstrap procedures. If inference is conducted according to the *misleading* asymptotic critical values, quite some evidence for cointegration is found, illustrated by the multitude of bold-starred entries in the two tables.

From the very strong evidence against cointegration across the variety of specifications we conclude that in many of the studies that use panel cointegration methods and asymptotic inference, the evidence for cointegration is mainly due to severely distorted small sample inference. This, of course, raises severe doubts on the validity of the results obtained in these studies.

In order to assess the potential mis-quantification of the BB and BS effect we will however also investigate equations based on 'cointegration'. The 'cointegrating' relationships are estimated by two methods, *fully modified* ordinary least squares (FM-OLS) and *dynamic* ordinary least squares (D-OLS). Both estimation methods are panel extensions of well known time series concepts. FM-OLS was introduced by Phillips and Hansen (1990) and D-OLS is due to Saikkonen (1991). Both methods allow for serial correlation in the residuals and for endogeneity of regressors in the cointegrating regression and result in asymptotically efficient estimation of the cointegrating vector. The panel extensions of FM-OLS are discussed in detail in Phillips and Moon (1999), nesting the discussion in Pedroni (2000) and Kao and Chiang (2000). As in the time series case the idea of FM-OLS is to obtain in the first step (OLS) estimates of long-run variance matrices. In the second step another regression is run on *corrected* variables, with the correction factors being functions of the estimated long-run variance matrices. The idea of D-OLS is to correct for serial correlation and endogeneity by augmenting the cointegrating regression by leads and lags of differences of the regressors. The panel extension of D-OLS is discussed in Mark and Sul (2001) and Kao and Chiang (2000). Both methods lead to asymptotically normally distributed (for both T and N to infinity) estimated cointegrating vectors, which implies that χ^2 inference via e.g. Wald tests can be conducted. Note for completeness that various implementations of both FM-OLS and D-OLS in a weighted or unweighted fashion are possible, see Hlouskova and Wagner (2004b) for a description. In this paper we do not discuss further details in this respect.

The last two columns of Table 4 display information concerning the results of FM-OLS and D-OLS estimation. A '+' indicates that in the estimation of the equation all coefficients have signs in line with the theoretical model, whereas a '–' indicates that at least one coefficient has a sign not in line with the model. It is remarkable that although there is no evidence for cointegration (when relying upon any of the bootstrapping procedures), for most of the equations, the coefficients are estimated with correct signs.

The sign predictions (noting again that they are opposite for the q- and q_2-equations) for the equations with the price variables as the dependent variables are as follows: A- and B-equations: $\beta_2 > 0$, C- and D-equations: $\beta_2 > 0, \beta_3 > 0$ and E-equations $\beta_2 > 0, \beta_3 < 0, \beta_4 > 0$. For the equations labelled Wages, LC-LPT and LC-LPN $\beta > 0$ and for BBE $\beta_1 > 0$ and $\beta_2 > 0$ are implied by the theoretical model.

The coefficient sign for the price index differential of tradables respectively the real exchange rate for tradables is expected to be positive in all equations. This indeed holds true for all equations.

6 Econometric Analysis III: Quantification of the Baumol-Bowen and Balassa-Samuelson Effects

In this section we now turn to a quantification of the Baumol-Bowen and the Balassa-Samuelson effects. The results of the preceding sections, namely the prevalence of unit root nonstationarity for many variables and almost no evidence for cointegration, implies that the equations will be formulated for *growth rates*.

However, we also estimate panel error correction versions of the equations, to assess the differences in the implied BB and BS effects between equations entirely in growth rates and equations incorporating nonstationary 'error correction' terms. The panel error correction equations contain lagged residuals of the corresponding cointegrating regressions, which are due the lack of evidence for cointegration very likely nonstationary. The differences in the estimated effects between equations without and with such error correction terms is our measure of the *bias* introduced by inappropriately resorting to cointegration techniques. For notational brevity throughout this section we refer to the equations that contain nonstationary (due to the absence of cointegration) error correction terms as specifications with cointegration. This is not meant to indicate cointegration!

In Table 4 the last two columns indicate for all equations, whether the signs of the coefficients in the 'cointegrating relationships' are in accordance with the theoretical predictions. As indicated in the previous section, two methods have been employed FM-OLS and D-OLS. It is remarkable that for most of the equations, both methods result in coefficient signs in line with theory. An exception to this observation are, however, three of the E-equations. The empirical results are based on the cointegration estimation method that results in the 'best' estimates. With 'best' here indicating that the coefficient signs of the coefficients in both the cointegrating relation and the resulting error correction equation are in line with the theory, i.e. of correct sign and significant.

The small sample size of our panels does not facilitate estimation, since many of the more advanced panel estimators like DPD (see e.g. Arellano, 2003) are known to perform poorly in small samples.[14] We thus proceed as follows in our estimation strategy for all equations,

[14]The optimality properties of these and related panel GMM estimators rest upon the cross-section dimension

both without and with error correction terms: We start by specifying equations by feasible GLS, where we allow for cross-section heteroscedasticity and correlation. The t-statistics are based on the feasible GLS specification. Potential endogeneity is of course a concern. Therefore we apply the Durbin-Wu-Hausman test for testing the null of consistency of GLS against the alternative of inconsistency due to regressor endogeneity (see e.g. Davidson and MacKinnon, 1993). This is done via auxiliary regressions, where in a first step the potentially endogenous regressors are each regressed on the set of instruments specified for the equation at hand. The residuals of these regressions are then added to the 'original' equation and the null hypothesis of all their coefficients being jointly equal to zero is tested. The advantage of the formulation of the Durbin-Wu-Hausman test via auxiliary regressions is to avoid to perform instrumental variables estimation already in the testing step. The precise details of potentially endogenous regressors and the corresponding instruments are available from the authors upon request. We perform the tests with various sets of regressors treated as potentially endogenous and various instrument sets. Usually the instruments are given by either lagged variables or variables for the EU11. The latter choice stems from the fact that many of the regressors are given by the difference of a variable in the CEE countries and the EU11, an example being $\Delta a_{it} - \Delta a_t^*$, for which Δa_t^* is an instrument candidate. Variables of this type are by construction good instruments: they are correlated with the variables and very likely uncorrelated with the error terms. In case of over-identification we have performed the corresponding tests for instrument validity, often referred to as J-test or Sargan test, see Arellano (2003). No rejections of the null occurred. All test results have to be seen in the light of the small sample, of course.

The results can briefly be summarized as follows: The null hypothesis of the Durbin-Wu-Hausman test was only rejected for equations ΔAp_2, ΔCp_{ec} and ΔCq_{ec}. Thus, only for these three equations 2SLS estimation is necessary. Only, for the equation ΔAp_2 a 2SLS specification with all coefficient signs in line with theory has been achieved (with the instruments given by the lagged price of tradables in the EU11 and the lagged BS variable). For the other two equations no specification with all coefficient signs according to theory has been obtained. Those two equations with error correction terms are thus not considered further.

It has been mentioned in the introduction that the BB and BS models offer purely supply side based explanations of price respectively real exchange rate movements. Since we are now

tending to infinity. Thus, for our sample with eight countries, no practical advantage can be expected.

	Δa^T	Δa^N	Δlcr^T		Δlcr^N		ΔlnY^T	ΔlnY^N
$(\frac{GFCF}{GDP})_{-1}$	-0.453	0.037						
	(-8.005)	(1.466)						
FDI	0.111	0.149						
	(1.294)	(1.283)						
Δa^T			0.772	0.403			0.868	
			(24.024)	(12.169)			(30.134)	
Δa^N					1.494	1.004		
					(18.288)	(55.919)		
Δa^N_{-1}								0.187
								(3.606)
ΔlnL^T_{-1}			-0.274	-0.090			0.156	
			(-5.675)	(-2.439)			(2.377)	
ΔlnL^N								0.598
								(11.009)
ΔlnL^N_{-1}					0.144	0.304		
					(1.647)	(12.952)		
$\frac{U}{U+L}$			-0.534					
			(-3.680)					
$\Delta lnGDPPC_{-1}$	0.047	0.024					0.401	
	(3.171)	(4.853)					(9.703)	
$\Delta lnTC_{-1}$								0.202
								(6.152)
$\hat{ec}_{-1}$				-0.630		-0.814		
				(-15.135)		(-27.088)		
$Adj.R^2$	0.055	0.063	0.399	0.558	0.432	0.632	0.847	0.373

Table 7: Testing for the presence of demand side effects on key variables in the BB and BS models. In the first column the dependent variables are listed and in the first row the regressors are listed. $\hat{ec}$ (*error correction*) denotes the residuals from the equation LC-LPT respectively LC-LPN, both estimated with FM-OLS.
Robust t-statistics are displayed in brackets. All equations include fixed effects.

focusing in our study on equations in growth rates, it becomes of particular importance to test for the influence of demand variables, which are potentially more important in the short and medium-run. We thus study next the potential impact of demand side variables on the evolution of the key variables in the BB and BS-models, see Table 7. The equations presented in the table are the result of extensive specification searches, where again also validity of GLS estimation has been verified by applying the Durbin-Wu-Hausman test as discussed. The analysis is inspired by Bergstrand (1991) and Halpern and Wyplosz (2002).

Let us start with a discussion of the equations for Δa^T and Δa^N. The premise concerning productivity developments in the BB and BS models is that productivity is supply driven. This is not fully confirmed by the two equations, where foreign direct investment is significant but also lagged real per capita GDP is significant.[15] Thus, demand side variables contribute to the evolution of labor productivities. Note furthermore the highly significant impact of foreign direct investment on productivity. This is consistent with the installment of efficient technologies by foreign investors via e.g. greenfield investments.

The next two columns correspond to the equations LC-LPT and LC-LPN already discussed in the previous section, now in growth rates and including additional explanatory variables. Also specifications including error correction terms are presented, the fourth and the sixth equation in Table 7. Productivity in the respective sectors is highly significant. However, also (for both sectors) employment growth is a further significant supply side variable, with negative sign in the tradables sector and with positive sign in the non-tradables sector. This supports the hypothesis that in transition productivity is driven i.a. by labor reallocation from the tradables (which is essentially equal to industry) to the non-tradables sector (which contains all services), see also Grafe and Wyplosz (1999). Thus, labor costs are driven by labor productivity and employment reallocation, i.e. labor costs are indeed supply side determined.

The final two equations assess the importance of supply and demand factors for output growth in both sectors. The supply side is captured by productivity and employment. However, also demand side variables are significant: Real per capita GDP in the tradables sector and total consumption in the non-tradables sector.

The above analysis shows that demand variables contribute to the explanation of productivity and output developments. Thus, in the subsequent econometric analysis the equations presented in levels in Table 4 have been estimated in growth rates and augmented by demand variables like real per capita GDP or total consumption. In Table 8 the resulting equations are listed. These represent again the results of extensive specification searches, in this case over explanatory demand side variables. For notational simplicity we do, however, not change the equation labels, except for a Δ to indicate the growth rate specification. Real per capita GDP is the main explanatory demand variable. This is perfectly consistent with Bergstrand (1991), who provides a corresponding extended theoretical model as well as an empirical study.

[15] Fischer (2002) presents an extension of the BB model where investment is affecting the (internal) real exchange rate. Note that in our case it is only foreign direct investment that is significant.

Label	Equation			SUR
ΔBBE	Δp_{it}^{rel}	$=$	$c_i + \beta_1 \Delta a_{it}^{rel} + \beta_2 \Delta w_{it}^{rel} + u_{it}$	$+$
ΔAq	Δq_{it}	$=$	$c + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(\delta_{it-1}(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel}) - \delta_{t-1}^*(a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + u_{it}$	$+$
ΔAq_2	$\Delta q_{2,it}$	$=$	$c + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(\delta_{it-1}(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel}) - \delta_{t-1}^*(a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + u_{it}$	$+$
ΔAp	$\Delta p_{it}^{GDP} - \Delta p_t^{GDP*}$	$=$	$c + \beta_1(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_2 \Delta(\delta_{it-1}(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel}) - \delta_{t-1}^*(a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + u_{it}$	$+$
ΔAp_2	$\Delta p_{it}^{T+N} - \Delta p_t^{(T+N)*}$	$=$	$c + \beta_1(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_2 \Delta(\delta_{it-1}(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel}) - \delta_{t-1}^*(a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + \beta_3 \Delta lnGDPPC_{it-1} + u_{it}$	$+$
ΔBq	Δq_{it}	$=$	$c_i + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel} - (a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + u_{it}$	$+$
ΔBq_2	$\Delta q_{2,it}$	$=$	$c_i + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel} - (a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + u_{it}$	$+$
ΔBp	$\Delta p_{it}^{GDP} - \Delta p_t^{GDP*}$	$=$	$c_i + \beta_1(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_2 \Delta(a_{it}^{rel} + \alpha_{it}^N w_{it}^{rel} - (a_t^{rel*} + \alpha_t^{N*} w_t^{rel*})) + u_{it}$	$+$
ΔBp_2	$\Delta p_{it}^{T+N} - \Delta p_t^{(T+N)*}$	$=$	$c + \beta_1(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_2 \Delta(a_{it-1}^{rel} + \alpha_{it-1}^N w_{it-1}^{rel} - (a_{t-1}^{rel*} + \alpha_{t-1}^{N*} w_{t-1}^{rel*})) + \beta_3 \Delta lnGDPPC_{it-1} + u_{it}$	$+$
ΔCq	Δq_{it}	$=$	$c_i + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 \Delta w_{BS,it-1}^{rel} + u_{it}$	$+$
ΔCq_2	$\Delta q_{2,it}$	$=$	$c_i + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 \Delta w_{BS,it}^{rel} + u_{it}$	$-$
ΔCp	$\Delta p_{it}^{GDP} - \Delta p_t^{GDP*}$	$=$	$c + \beta_1(\Delta p_{it-1}^{GDP} - \Delta p_{t-1}^{GDP*}) + \beta_2(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_3 \Delta(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_4 \Delta w_{BS,it}^{rel} + \beta_5 \Delta lnGDPPC_{it-1} + u_{it}$	$+$
ΔCp_2	$\Delta p_{it}^{T+N} - \Delta p_t^{(T+N)*}$	$=$	$c_i + \beta_1(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_2 \Delta(\delta_{it} a_{it}^{rel} - \delta_t^* a_t^{rel*}) + \beta_3 \Delta w_{BS,it}^{rel} + \beta_4 \Delta lnGDPPC_{it-1} + u_{it}$	$+$
ΔDq	Δq_{it}	$=$	$c_i + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(a_{it}^{rel} - a_t^{rel*}) + \beta_3 \Delta w_{BS,it-1}^{rel} + u_{it}$	$+$
ΔDq_2	$\Delta q_{2,it}$	$=$	$c + \beta_1 \Delta q_{it}^T + \beta_2 \Delta(a_{it}^{rel} - a_t^{rel*}) + \beta_3 \Delta lnTC_{it} + u_{it}$	$+$
ΔDp	$\Delta p_{it}^{GDP} - \Delta p_t^{GDP*}$	$=$	$c + \beta_1(\Delta p_{it-1}^{GDP} - \Delta p_{t-1}^{GDP*}) + \beta_2(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_3 \Delta(a_{it}^{rel} - a_t^{rel*}) + \beta_4 \Delta w_{BS,it}^{rel} + \beta_5 \Delta lnGDPPC_{it-1} + u_{it}$	$+$
ΔDp_2	$\Delta p_{it}^{T+N} - \Delta p_t^{(T+N)*}$	$=$	$c_i + \beta_1(\Delta p_{it-1}^{T+N} - \Delta p_{t-1}^{(T+N)*}) + \beta_2(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_3 \Delta(a_{it}^{rel} - a_t^{rel*}) + \beta_4 \Delta w_{BS,it}^{rel} + u_{it}$	$+$
ΔEp	$\Delta p_{it}^{GDP} - \Delta p_t^{GDP*}$	$=$	$c_i + \beta_1(\Delta p_{it}^T - \Delta p_t^{T*}) + \beta_2 \Delta a_{it}^{rel} + \beta_3 \Delta a_t^{rel*} + \beta_4 \Delta w_{BS,it}^{rel} + u_{it}$	$-$

Table 8: Baumol-Bowen and Balassa-Samuelson equations in growth rates extended by demand side variables.
A '+' in the last column indicates that all coefficient signs are in line with the theoretical model, '-' indicates that at least one coefficient sign is not according to the model.
Equations ΔE_q, ΔE_{q2} and ΔE_{p2} (cf. Table 4) are dropped, since for them the null hypothesis of the coefficient of Δa_{it}^{rel} being equal to minus the coefficient to Δa_t^{rel*} could not be rejected. Thus, for these the corresponding (restricted) D-equations are reported.

We have noted previously that the E-equations nest the D-equations. The corresponding test, performed for the four different ΔE-equations is given by the null hypothesis $\beta_2 = -\beta_3$, with β_2 denoting the coefficient to relative labor productivity in the CEE countries and β_3 denoting the coefficient corresponding to relative labor productivity in the EU11. For three out of the four equations the null hypothesis is not rejected. Thus, in Table 8 the equations ΔDq, ΔDq_2, ΔDp_2 and ΔEp are displayed. The last column in Table 8 displays again the information concerning the coefficient signs. Here again a '+' indicates that the signs of all estimated coefficients are in line with the theoretical predictions, which hold for all but two equations (ΔCq_2 and ΔEp). These latter two equations, ΔCq_2 and ΔEp, will thus not be considered further in the quantification of the BS effects and the subsequent inflation simulations presented in Section 7. Albeit ΔDp_{2ec} has one coefficient with wrong sign (corresponding to total consumption) we keep it, as only one coefficient sign is incorrect.

Throughout, the equations with error correction terms are presented with the same names as the corresponding equations without error correction terms with a further subscript ec added. We report only those equations with error correction terms where both the coefficients of the estimated equation (in growth rates) and the coefficients in the 'cointegrating relationships' all have correct signs (nine in total).

The results of the estimations are presented in two tables: In Table 9 we report the equations with the real exchange rates (Δq and Δq_2) as dependent variables and in Table 10 the equations with the inflation differentials as dependent variables are displayed. In that table also the extended Baumol-Bowen equation BBE with Δp^{rel} as dependent variable is included. The equations are then used below to quantify the Baumol-Bowen respectively the Balassa-Samuelson effect and also for projections concerning the evolution of the inflation rates in Section 7.

	ΔAq	ΔAq_{ec}	ΔAq_2	ΔAq_{2ec}	ΔBq	ΔBq_{ec}	ΔBq_2	ΔBq_{2ec}	ΔCq	ΔDq	ΔDq_2	ΔDq_{2ec}
Interc.	-0.037	-0.035	-0.041	-0.040	FE	FE	FE	FE	FE	FE	-0.020	-0.028
	(-11.991)	(-12.267)	(-12.028)	(-6.017)							(-5.084)	(-20.454)
Δq^T	0.662	0.659	0.637	0.602	0.668	0.739	0.627	0.650	0.587	0.672	0.815	0.717
	(26.807)	(34.015)	(24.429)	(10.802)	(52.987)	(10.606)	(30.166)	(31.874)	(24.533)	(38.229)	(25.288)	(101.315)
ΔBS									-0.304			
									(-5.981)			
$\Delta BSE1_{-1}$	-0.169	-0.023	-0.175	-0.074								
	(-6.554)	(-1.186)	(-6.860)	(-1.306)								
$\Delta BSE2_{-1}$					-0.082	-0.031	-0.097	-0.007				
					(-17.079)	(-0.913)	(-14.374)	(-0.854)				
$\Delta(a^{rel} - a^{rel*})$										-0.046	-0.049	-0.139
										(-2.336)	(-1.456)	(-14.230)
$\Delta w^{rel}_{BS,-1}$									-0.103	-0.104		
									(-5.217)	(-7.401)		
$\Delta lnTC$											-0.280	-0.171
											(-4.925)	(-18.581)
$\hat{ec}_{-1}$		-0.353		-0.237		-0.058		-0.085				-0.159
		(-12.144)		(-4.680)		(-2.926)		(-14.361)				(-22.172)
Adj.R^2	0.679	0.797	0.704	0.783	0.685	0.733	0.718	0.788	0.687	0.662	0.748	0.706

Table 9: Estimation results for equations in growth rates with the real exchange rate variables as dependent variables. FE in the row labelled *Interc.* indicates that fixed effects are included. For the other equations the null hypothesis of equal fixed effects could not be rejected and therefore a common intercept is estimated.
The regressor $\hat{ec}$ denotes the 'cointegrating' relationship from the corresponding equation in levels as displayed in Table 4.
In brackets robust t-statistics are displayed.

We first discuss the results in Table 9. The equation ΔCq_2 has been excluded, since no specification with all coefficient signs in line with theory could be obtained, as already mentioned above. For equations ΔCq and ΔDq no error correction specifications with all coefficient signs in line with theory could be obtained. For all other equations, both a specification in growth rates and as error correction model, with all coefficient signs in line with theory has been found.

A couple of important observations can be made concerning the final specifications: In the ΔA- and ΔB-equations the BS variables, $\Delta BSE1$ and $\Delta BSE2$ enter lagged, as does the relative wage variable Δw_{BS}^{rel} in equations ΔCq and ΔDq. Furthermore, and this is a difference to the equations in the next table, almost nowhere are demand variables significant. Only total consumption is significant in two of the D-equations. The error correction terms are obtained by D-OLS for equations ΔBq_{ec} and ΔDq_{2ec} and by FM-OLS for the other equations. Note that the adjusted R^2 is with one exception (the A-equation with error correction term) higher for the equations with Δq_2 as the dependent variable. This is not surprising, since the theoretical model is specified for the tradables and non-tradables sectors only. Thus, we expect better fit for a corresponding dependent variable. This is confirmed by the results. The (rate of change of the) real exchange rate of tradables is highly significant throughout.

Next we turn briefly to the equations displayed in Table 10. Error correction specifications with all coefficient signs correct have been found for ΔBBE, ΔBp, ΔDp and ΔDp_2. Note again that in the equation ΔDp_2 with the error correction term total consumption enters with the wrong sign. Also note the large t-values for the equation ΔDp_{ec}. These stem from the inclusion of the nonstationary error correction term. Spurious regression often manifests itself in large t-values, and it is actually surprising that this effect shows up only in one of the equations with error correction terms. For all equations the error correction term is estimated by FM-OLS. For many of the equations again the lagged BS variables remains after specification analysis. Also, lagged real per capita GDP or lagged total consumption growth stay in the final specification in several cases. The prevalence of lagged variables indicates a certain degree of stickiness in the transmission mechanism outlined by the BB and BS models. The inflation differential in tradables between the CEE countries and the EU11 is highly significant in all equations. As indicated above, only for the equation ΔAp_2 2SLS estimation is warranted by the Durbin-Wu-Hausman test.

	ΔBBE	ΔBBE_{ec}	ΔAp	ΔAp_2	ΔBp	ΔBp_{ec}	ΔBp_2	ΔCp	ΔCp_2	ΔDp	ΔDp_{ec}	ΔDp_2	ΔDp_{2ec}
Interc.	FE	FE	0.025	FE	0.021	0.043	0.016	FE	FE	0.013	0.014	FE	FE
			(5.291)		(5.178)	(12.226)	(2.739)			(4.686)	(66.242)		
$\Delta p^T - \Delta p^{T*}$			0.869	1.096	0.957	0.619	0.960	0.821	0.897	0.997	1.044	1.096	0.787
			(32.383)	(11.855)	(47.990)	(17.626)	(31.385)	(62.696)	(58.781)	(56.775)	(764.545)	(70.774)	(31.017)
ΔBS									0.222				
									(4.342)				
ΔBS_{-1}								0.143					
								(4.471)					
$\Delta BSE1_{-1}$			0.081	0.138									
			(3.221)	(2.570)									
$\Delta BSE2$					0.013	0.032							
					(0.963)	(3.751)							
$\Delta BSE2_{-1}$							0.034						
							(2.932)						
$\Delta(a^{rel} - a^{rel*})$										0.229	0.170	0.255	0.105
										(6.525)	(118.318)	(7.519)	(4.775)
Δa^{rel}	0.229	0.297											
	(4.098)	(2.654)											
Δw^{rel}	0.398	0.154											
	(4.688)	(1.809)											
Δw^{rel}_{BS}									0.062	0.069	0.118	0.192	0.140
									(2.707)	(1.981)	(91.977)	(5.937)	(7.106)
$\Delta lnGDPPC_{-1}$							0.134		0.059				
							(1.962)		(1.423)				
$\Delta lnTC$								0.242					
								(4.507)					
$\Delta lnTC_{-1}$												0.287	-0.075
												(6.281)	(-1.521)
$\hat{ec}_{-1}$		-0.566				-0.239					-0.234		-0.470
		(-6.510)				(-14.017)					(-80.345)		(-7.938)
$Adj.R^2$	0.057	0.446	0.922	0.901	0.978	0.811	0.748	0.924	0.942	0.980	0.725	0.975	0.793

Table 10: Estimation results for equations in growth rates with the inflation differentials as dependent variables. FE in the row labelled *Interc.* indicates that fixed effects are included. For the other equations the null hypothesis of equal fixed effects could not be rejected and therefore a common intercept is estimated.
The regressor $\hat{ec}$ denotes the 'cointegrating' relationship from the corresponding equation in levels as displayed in Table 4.
In brackets robust t-statistics are displayed.

The above equations in Tables 9 and 10 form the basis for an assessment of the Baumol-Bowen and the Balassa-Samuelson effects. The quantification is given by the product of the estimated coefficient corresponding to the Baumol-Bowen respectively Balassa-Samuelson variable times the average value of the BB or BS-variable. Again, two periods are considered, 1994–2001 and 2000–2001.

We present three different estimates of the effect. One based on the specifications without error correction terms and two based on the specifications with error correction terms. Since none of the equations contains the lagged dependent variable as a regressor and either the contemporaneous BS variable or a certain lag of the BS variable is included only, no distinction has to be made between short and long-run effects.[16] In the specifications including the error correction term short- and long-run effects are to be distinguished. The short-run effect is given by the estimated coefficient corresponding to the BB or BS variable times the average value of the variable over the period considered. The long-run effect with error correction is derived entirely from the error correction term. It is given by the product of the coefficient corresponding to the BB or BS variable in the 'cointegrating relationship' times the average value of the variable. This uses the well known fact that in cointegrating regression the long-run elasticity is given by the corresponding coefficient in the cointegrating relationship.

We start with a discussion of the Baumol-Bowen effect, summarized in Table 11. The average productivity growth rates in both the tradables and the non-tradables sectors for the two periods considered are displayed in Table 1. The negative estimated Baumol-Bowen effects for the period 2000–2001 for the Czech Republic, Latvia and the Slovak Republic directly follow from the fact that over these two years productivity growth in these countries is higher in the non-tradables sector than in the tradables sector. For the other countries, and for all countries for the longer period, as discussed already in Section 3, productivity growth is higher in the tradables sector than in the non-tradables sector. The results are quite clear: The effect in the equation without cointegration is for all countries smaller than both measures of the effect derived from the equation containing the error correction term. Over the larger period the dual inflation rate contribution ranges from 0.15% for Estonia to 1.36% for Hungary. It is probably noteworthy that with the exception of the Baltic countries, 'similar' countries' estimate of the BB effect are rather similar. Within the Baltic countries

[16]Note that this is, of course, just the usual distinction in econometrics between short- and long-run elasticities. Due to the result of the specification search, no distinction has to be made for the equations in growth rates, upon which we will later on base our inflation simulations.

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN
				1994–2001				
no coint.	0.682	0.151	1.360	0.197	0.925	1.219	0.348	1.100
SR with coint.	0.887	0.196	1.768	0.256	1.202	1.585	0.452	1.430
LR with coint.	2.522	0.558	5.027	0.728	3.417	4.506	1.287	4.067
				2000–2001				
no coint.	-0.518	0.298	0.723	-0.572	1.596	0.855	-0.291	1.117
SR with coint.	-0.674	0.387	0.940	-0.744	2.074	1.112	-0.378	1.452
LR with coint.	-1.916	1.102	2.674	-2.115	5.898	3.161	-1.076	4.130

Table 11: Estimates of the Baumol-Bowen effect for the CEEC8 in percent of dual inflation per year. The Baumol-Bowen effect is given by the product of the coefficient to Δa^{rel} and the average value of this variable over the indicated period.
The first lines labelled *no coint.* display the effect based on the equation ΔBBE and the second and third lines display the short- and long-run effects based on the equation ΔBBE_{ec}.

Lithuania is sort of an outlier. This might be due to the different exchange rate regime. Both, Estonia and Latvia operate currency boards, whereas Lithuania follows a fixed peg strategy.

The effects of adding a cointegration term are substantial. As already mentioned, both the short and long-run effects are bigger for all countries than the effects estimated from the equations without error correction terms. In particular the long-run effects are bigger by a factor of about four for all countries.

Similar observations as for the longer period can also be made for the shorter period, with the 'correct' no cointegration estimates now ranging from -0.57% for Latvia to 1.60% for Lithuania. Adding error correction terms again increases the effect, as for the longer period by approximately a factor four when deriving the (long-run) effect from the cointegrating regression.

In Table 12 we display the period averages for the Balassa-Samuelson terms for equations ΔA, with $BSE1$, to ΔD, with $\Delta a^{rel} - \Delta a^{rel*}$ as BS variable. The different terms are, as already discussed in Section 2, based on equations (8) to (11). The two variables most closely related to the Cobb-Douglas specification of the theoretical model are $\Delta BSE1$ and $\Delta BSE2$. Both variables are negative for all CEE countries for both periods considered. Thus, the estimated BS effect for equations ΔA and ΔB will be *negative* for all countries, since in the specifications *correct* coefficient signs are prevalent throughout. Comparing the BS-terms $BSE1$ and $BSE2$ with the other BS terms for equations ΔC and ΔD, it is seen that the main difference is the inclusion of the relative wage terms. For comparison, in Figure 2 where

	$\Delta BSE1$	$\Delta BSE2$	ΔBS	$\Delta a^{rel} - \Delta a^{rel*}$
Averages over 1994–2001				
CZE	-1.846	-0.445	-0.391	1.153
EST	0.217	-5.929	1.704	-1.169
HUN	-6.533	-5.288	-0.649	4.115
LVA	-0.230	-13.254	3.742	-0.969
LTU	-3.914	-11.224	2.160	2.212
POL	-3.825	-4.208	0.549	3.499
SVK	-4.041	-3.367	-1.634	-0.308
SVN	-6.426	-9.748	0.047	2.980
Averages over 2000–2001				
CZE	-3.651	-7.497	-1.391	-3.816
EST	-12.477	-25.432	0.529	-0.247
HUN	-3.169	-1.648	-0.229	1.612
LVA	-8.928	-24.226	0.559	-4.052
LTU	-12.637	-13.722	-1.233	5.425
POL	-3.891	-3.366	0.085	2.188
SVK	-9.922	-13.137	-3.367	-2.822
SVN	-20.191	-29.050	-2.664	3.334

Table 12: Average values of Balassa-Samuelson variables for equations ΔA to ΔD.

on the horizontal axis $\Delta a^{rel} - \Delta a^{rel*}$ is displayed (i.e. a BS-variable that does not include the relative wage terms), a negative BS effect is only prevalent for three or four countries. The difference stems from the intermingling of the contributions of relative productivity growth and relative wage growth in $BSE1$ and $BSE2$. In the quantification of the BS-effect below we want to separate these two parts and thus focus on the ΔC and ΔD equations. Note, however, that the estimated effect (being the product of the BS variable and the corresponding coefficient) *do depend* upon the relative wages, since we only consider equations that include relative wage terms as regressors because of wage non-homogeneity.

The BS variables corresponding to equations ΔC and ΔD, namely ΔBS and $\Delta a^{rel} - \Delta a^{rel*}$ are displayed in the last two columns of Table 12. Especially the latter is a common choice in the literature. Both of these variables are positive for a majority of countries for the longer period and for about half of the countries for the shorter period. Note that the set of countries for which the variables, in particular $\Delta a^{rel} - \Delta a^{rel*}$, are negative for the shorter period, include the three countries for which the Baumol-Bowen effect was found to be negative above.

For the ΔC- and ΔD-equations we now discuss the three different measures of the Balassa-Samuelson effect, similarly to the above quantification of the Baumol-Bowen effect. We start

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN
Quantification of effect with Δq as dependent variable								
ΔCq, 1994–2001								
no coint.	0.119	-0.518	0.197	-1.137	-0.656	-0.167	0.497	-0.014
ΔCq, 2000–2001								
no coint.	0.423	-0.161	0.070	-0.170	0.375	-0.026	1.023	0.809
ΔDq, 1994–2001								
no coint.	-0.053	0.054	-0.190	0.045	-0.102	-0.162	0.014	-0.138
ΔDq, 2000–2001								
no coint.	0.177	0.011	-0.075	0.188	-0.251	-0.101	0.131	-0.154
Quantification of effect with Δq_2 as dependent variable								
ΔDq_2, 1994–2001								
no coint.	-0.056	0.057	-0.201	0.047	-0.108	-0.171	0.015	-0.146
SR with coint.	-0.160	0.162	-0.570	0.134	-0.306	-0.485	0.043	-0.413
LR with coint.	-1.815	1.840	-6.474	1.524	-3.480	-5.506	0.484	-4.689
ΔDq_2, 2000–2001								
no coint.	0.187	0.012	-0.079	0.198	-0.265	-0.107	0.138	-0.163
SR with coint.	0.529	0.034	-0.223	0.561	-0.752	-0.303	0.391	-0.462
LR with coint.	6.003	0.389	-2.536	6.375	-8.536	-3.443	4.441	-5.245

Table 13: Balassa-Samuelson effect in % in equations for (rate of change of) real exchange rate measures. The Balassa-Samuelson effect is defined as the product of the coefficient to the BS-variable in the corresponding equations with the average values of the variables as displayed in Table 12.
no coint. indicates the effect from the specification without the error correction term, *SR with coint.* indicates the short-run effect derived from the specification with the error correction term and *LR with coint.* indicates the long-run effect derived from the cointegrating regression in the corresponding equation.

in Table 13 with the quantification in the equations for the real exchange rate measures and after that turn in Table 14 to the equations with the inflation differentials as dependent variables.

In Table 13 the upper panel displays the effect when the rate of change of the real exchange rate, Δq, is the dependent variable. The following main observations emerge: Both the ordering across countries and the magnitude of the estimated BS effect differ between the ΔC-and the ΔD-equation. The ordering across countries for the ΔD-equations is the same as the one for the Baumol-Bowen effect. This holds also when using Δq_2 instead of Δq as dependent variable. Note here that the contribution of the BS effect to the real exchange rate evolution is an *appreciation* of the real exchange rate, i.e. a *decline* of q. The magnitude of the effect varies substantially between the two periods 1994–2001 or 2000–2001, with a general tendency to be smaller over the shorter period. The largest appreciation is about 1.14% for Latvia based on the ΔC-equation and the period 1994–2001. Basing the effect on only the years 2000–2001 implies for all countries a smaller rate of appreciation, or – for the countries with *depreciation* except for Hungary – larger depreciation for the ΔC-equation. For the ΔD-equation there is no unidirectional change between the effects computed from the 1994–2001 averages to the effects computed from the 2000–2001 averages. Furthermore, the effect based on the ΔD-equation is smaller than the effect based on the ΔC-equation for all countries but Slovenia. As with the BB effect, resorting to cointegrating equations leads to an effect that is bigger by a factor four on average.

Using Δq_2 as the dependent variable, shown in the lower panel of Table 13, results on average in smaller appreciation, respectively depreciation than using Δq. For the shorter period 2000–2001 the difference between the assessment without and with the 'error correction' term is huge (up to a factor 30) for all countries but Estonia. For the ΔD-equation a bigger effect is found for a majority of countries when computed from the smaller period averages.

For the equations with the inflation rate differentials as dependent variables the following observations can be made in Table 14. The ranking of the effect across countries is the same for both dependent variables for all ΔC specifications and for all ΔD specifications for each of the two periods. Note, however, that the rankings differ between the ΔC and ΔD equations and also between the periods. The details are as follows: For ΔCp and ΔCp_2 over the period 1994–2001 the ranking, based on the equations without cointegration terms, is (from largest to smallest inflation differential) Latvia, Lithuania, Estonia, Poland, Slovenia, Czech Republic, Hungary and Slovak Republic. Thus, the three Baltic countries have the largest BS-effect according to the ΔC-equations. For ΔDp and ΔDp_2 and the same period the corresponding ranking is Hungary, Poland, Slovenia, Lithuania, Czech Republic, Slovak Republic, Latvia and Estonia.

Looking in more detail at the effect obtained from the equations with the GDP deflator inflation differentials, the following observations emerge. The differences between the effects without and with cointegration are less pronounced than in the case of the equations for the real exchange rates, although still in general the use of cointegration results in larger effects. For the shorter period the effect estimated given by the long-run effect of the cointegrating relation leads to effects that are larger by a factor two to three than the effects without cointegration. Basing the effect on values for 1994–2001, the largest inflation differentials are

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN
Quantification of effect with $\Delta p - \Delta p^*$ as dependent variable								
ΔCp, 1994–2001								
no coint.	-0.056	0.243	-0.093	0.534	0.308	0.078	-0.233	0.007
ΔCp, 2000–2001								
no coint.	-0.198	0.075	-0.033	0.080	-0.176	0.012	-0.480	-0.380
ΔDp, 1994–2001								
no coint.	0.264	-0.267	0.941	-0.222	0.506	0.801	-0.070	0.682
SR with coint.	0.196	-0.199	0.700	-0.165	0.376	0.595	-0.052	0.507
LR with coint.	0.795	-0.806	2.837	-0.668	1.525	2.413	-0.212	2.055
ΔDp, 2000–2001								
no coint.	-0.873	-0.057	0.369	-0.927	1.241	0.501	-0.646	0.763
SR with coint.	-0.649	-0.042	0.024	-0.689	0.923	0.372	-0.480	0.567
LR with coint.	-2.631	-0.171	1.111	-2.794	3.741	1.509	1.946	2.299
Quantification of effect with $\Delta p_2 - \Delta p_2^*$ as dependent variable								
ΔCp_2, 1994–2001								
no coint.	-0.087	0.378	-0.144	0.829	0.479	0.122	-0.362	0.011
ΔCp_2, 2000–2001								
no coint.	-0.308	0.117	-0.051	0.124	-0.273	0.019	-0.746	-0.590
ΔDp_2, 1994–2001								
no coint.	0.294	-0.298	1.050	-0.247	0.564	0.893	-0.079	0.760
SR with coint.	0.122	-0.123	0.434	-0.102	0.233	0.364	-0.032	0.314
LR with coint.	0.337	-0.341	1.201	-0.283	0.646	1.022	-0.090	0.870
ΔDp_2, 2000–2001								
no coint.	-0.973	-0.063	0.411	-1.034	1.384	0.558	-0.720	0.850
SR with coint.	-0.402	-0.026	0.170	-0.427	0.572	0.231	-0.298	0.352
LR with coint.	-1.111	-0.072	0.470	-1.183	1.584	0.639	-0.824	0.973

Table 14: Balassa-Samuelson effect in % in equations for the inflation differentials. The Balassa-Samuelson effect is defined as the product of the coefficient to the BS-variable in the corresponding equations with the average values of the variables as displayed in Table 12.
no coint. indicates the effect from the specification without the error correction term, *SR with coint.* indicates the short-run effect derived from the specification with the error correction term and *LR with coint.* indicates the long-run effect derived from the cointegrating regression in the corresponding equation.

found for Hungary with 0.94% and Poland with 0.80% and the smallest are found for Estonia with -0.27% and the Slovak Republic with -0.23%. For most countries for both periods the 'correct' (i.e. without cointegration) BS-effect is smaller than about half a percent.

When the inflation differential is only computed with respect to the prices in the two sectors tradables and non-tradables, see the lower panel of Table 14, qualitatively the same observations as for the GDP deflator based inflation rates emerge. A larger effect is obtained when using cointegration, where partly the differences are rather large again between the specifications without and with cointegration. The largest BS effect is again observed for Hungary, now with 1.05%. Again, the effects are, with few exceptions, smaller than half a percent.

The results discussed above show that the BB or BS effect alone are not very powerful in explaining the evolution of the real exchange rate respectively the inflation differentials between the CEECs and the EU11. The effects are, as we have seen, mostly below half a percent and partly even negative. This is, however, not too surprising, given the fact that several key assumptions of the standard models are not fulfilled. These are wage homogeneity, PPP in tradables and the irrelevance of demand side factors. It is only the inclusion of these additional explanatory variables that leads to well specified equations with significant coefficients with correct signs. In principle this has already been seen in Figure 2, where positive correlation between the productivity growth and inflation differentials is visible, but where one can also see that the degree of determination of this correlation is rather low. Thus, in order to assess the implications for the real exchange rates or inflation rates, which we focus upon, the impact of these variables has to be taken into account. In other words, the magnitude of the *pure* BB or BS effect is certainly not the best indicator of the explanatory power of the underlying models for the evolution of the real exchange rate respectively the inflation differentials. In the following section we therefore derive inflation projections based on not only the BB or BS terms but include also the other explanatory variables. This, of course, requires to make assumptions concerning all the explanatory variables.

7 Inflation Simulations

In this section we present the inflation simulations stemming from the analysis in the previous sections. Two sets of simulations are performed. One based on the Baumol-Bowen equation, discussed in subsection 7.1, and one based on the Balassa-Samuelson equations, discussed in subsection 7.2. The first set of simulations based on the ΔBBE equation is inspired by the simulations performed in Alberola and Tyrväinen (1998).[17] We perform inflation simulations for *all* BS-equations without the incorrect cointegration term. Performing simulations for all equations, including the ΔA to ΔB equations, reflects again the previously made observation that all the specified equations have all coefficients significant and with signs according to theory and fit the data rather well.[18] It is the inclusion of additional variables, like PPP-deviations, relative wages or demand variables that contributes importantly to the explanation of the dependent variable. In other words, it is not only the BS-term that is relevant for the inflation developments.

7.1 Baumol-Bowen Inflation Simulations

Let us start with a discussion of the Baumol-Bowen based inflation simulation. Following Alberola and Tyrväinen (1998) the inflation rate in the tradables sector is assumed identical for all countries. This allows to compute the country specific inflation rates in non-tradables, based on an assumption for aggregate inflation in the CEEC8 group. Therefore, to obtain a simulation for the GDP deflator based inflation rate requires to furthermore specify assumptions concerning inflation in agriculture and the public sector.

Denote with ρ_i the output share of country i in the group CEEC8. Then, the inflation rate in the group CEEC8 can be written as

$$\Delta p_{CEEC8} = \sum_{i=1}^{8} \rho_i \Delta p_i \tag{13}$$

where Δp_i is the GDP deflator inflation rate in country i. For notational simplicity the superscript GDP is omitted throughout in this section. Since the economy consists of four sectors and the ΔBBE equations is only concerned with the tradables and non-tradables sectors, one further step is necessary. The GDP deflator inflation rate is given by the weighted average of the inflation in the tradables and non-tradables sector, Δp_i^{T+N}, and by inflation in agriculture and the public sector, Δp_i^{A+P}. The weights are given by the respective GDP

[17] Simulations paralleling the discussion in Sinn and Reutter (2001), who ask, translated to our investigation, the question what minimum inflation rate is required in an enlarged monetary union in order to prevent deflation in any member state are also possible. To do so, however, would require a disaggregated analysis also of the EU11 countries. A detailed investigation of the BS-effect in the EU11 is performed in Wagner and Doytchinov (2004).

[18] The exception being equation Δp_{2ec} with the wrong sign for lagged total consumption.

shares, θ_i say.

$$\Delta p_i = \theta_i^{T+N} \Delta p_i^{T+N} + (1 - \theta_i^{T+N}) \Delta p_i^{A+P} \tag{14}$$

From the equation ΔBBE, the following representation for the inflation rate in the tradables and non-tradables sector together can be obtained:

$$\begin{aligned} \Delta p_i^{T+N} &= (1-\delta_i)\Delta p_i^T + \delta_i \Delta p_i^N && (15) \\ &= \Delta p_i^T + \delta_i \Delta p_i^{rel} && (16) \\ &= \Delta p_i^T + \delta_i(\hat{c}_i + \hat{\beta}_1 \Delta a_i^{rel} + \hat{\beta}_2 \Delta w_i^{rel}) && (17) \end{aligned}$$

For the inflation simulation it is seen from equation (17) that assumptions concerning the relative productivity growth and relative wage growth are required. We use the historical averages of Δa_i^{rel} and Δw_i^{rel}over three periods, 1994–2001, 1996–2001 and 2000–2001, see Table 32 in Appendix B. For the inflation simulation we also use the historical averages for the inflation rates in agriculture and the public sector over the same periods. These three periods, chosen according to the disinflation progress made in the CEEC8, see Table 31 in Appendix B, show the impact of different periods for the scenario variables on the inflation simulations.[19]

Equations (14) and (17) can be combined with an assumption on inflation in the CEEC8 (via equation (13)) and the assumption of equal tradable price inflation in all countries, to compute Δp^T. Then, for the computed value of Δp^T, inserting in equation (17) gives the implied inflation rate for country i in the tradables and non-tradables sector together, Δp_i^{T+N}. This value can now be combined with the assumed inflation in agriculture and the public sector for country i, Δp_i^{A+P}, to obtain the implied inflation for country i according to equation (14). We perform this simulation with two assumptions concerning inflation in the CEEC8 as a group of countries. The first assumption is $\Delta p_{CEEC8} = 2\%$. This assumption corresponds to the value that is often specified as at least an implicit inflation objective for the Euro Area but also other Western European countries. The results of this simulation are displayed in Table 15. The last column of this table displays the implied inflation rate for tradables, which is negative with values between 4 and 5% deflation. This, basically says that, an inflation objective of 2% annual inflation is only sustainable with substantial

[19] The average annual aggregate inflation rate in the CEEC8 was given by 12.25% over the period 1994–2001, by 8.98% over the period 1996–2001 and by 5.59% over the period 2000–2001. The year 1996 was the first year where the aggregate inflation rate in the CEEC8 was below 20%, at 14.62%.

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN	Δp^T_{CEEC8}
1994–2001	-2.31	1.52	2.14	2.08	-1.87	4.49	-1.15	0.07	-4.88
1996–2001	-2.55	0.11	2.21	1.92	-3.40	4.61	-1.11	0.04	-4.20
2000–2001	-2.48	-2.62	0.32	0.73	-4.65	5.48	-2.18	-0.17	-4.76

Table 15: Baumol-Bowen inflation simulations under the assumption of an aggregate inflation in the CEEC8 of 2% per annum.

deflation in tradables, under the assumption that inflation in agriculture and the public sector continues at the historical average values. This latter assumption is probably not too bad, given the fact that structural reforms in agriculture and the public sector, including abolishing price regulations, are essentially inevitable due to the EU membership of all countries in our sample. The simulation exercise results in deflation for the Czech Republic, Lithuania, the Slovak Republic and with the 2000–2001 values also for Estonia and Slovenia. This, of course, is rather unlikely. We thus, conclude from the BB simulation that 2% is too tight a target for the group. Note furthermore that the inflation predictions show substantial differences across countries.

The above results lead us to consider also the following experiment, with the results displayed in Table 16. Here we compute the inflation objective according to 2% inflation in the tradables and non-tradables sector for each country and to this we add the actual (over the corresponding period) inflation rates in agriculture and the public sector, using equation (14) for each country. Now, the inflation rate for the group CEEC8 is between 4 and 5%, see the last column in Table 16. Tradables inflation for the CEEC8 computed from these assumptions is -1.37 % (1994–2001), -1.55% (1996–2001) and -2.09% (2000–2001). Thus, still deflation in tradables prices results from the experiment. The implied inflation rates for the individual countries vary (depending upon the period used for the explanatory variables) between -2.67% for Lithuania to 7.60% for Poland. Still, some countries face deflation, based on this less restrictive Baumol-Bowen inflation simulation.

7.2 Balassa-Samuelson Inflations Simulations

The 15 (out of 17) equations with all coefficients signs according to theory displayed in Table 8 form the basis for the Balassa-Samuelson inflation simulations. We consider all well specified equations ΔA to ΔD.

Based on assumptions concerning the inflation rate in the EU11 and assumptions con-

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN	Δp_{CEEC8}
1994–2001	0.57	4.11	4.73	4.68	0.68	7.20	1.60	2.73	4.71
1996–2001	-0.37	2.10	4.19	3.92	-1.47	6.67	0.96	2.07	4.07
2000–2001	-0.26	-0.57	2.35	2.77	-2.67	7.60	-0.11	1.87	4.10

Table 16: Baumol-Bowen inflation simulations under the assumption of an aggregate inflation in the CEEC8 in tradables and non-tradables only of 2% per annum. The implied inflation rate for the CEEC8 country group is displayed in the last column.

cerning the explanatory variables in the equations, similar as above, inflation rates for the CEE countries can be computed. Care has to be taken of the fact that there are four different dependent variables in the equations. These imply slight differences for the computation of the inflation rates Δp_i. To assess the implied inflation rates from the equations with the real exchange rate measures as dependent variables, requires an assumption concerning the evolution of the nominal exchange rate of the CEEC's currency with respect to the Euro, since $\Delta q_i = \Delta e_i + \Delta p^* - \Delta p_i$. In our inflation simulations we assume that the nominal exchange rate does not change. Any assumed appreciation (depreciation) would reduce (increase) the implied inflation rate for country i. For the inflation rate in the EU11 we assume 2%.

Note also that for the equations with p_2 and q_2 as dependent variables, an assumption concerning inflation only in tradables and non-tradables in the EU11 is required, which we again set to 2%. Furthermore, of course, from here again equation (14) has to be invoked to compute the inflation rate for the GDP deflator. As for the above Baumol-Bowen experiment, the historical averages over the corresponding periods for inflation in agriculture and the public sector are used. The values used for the explanatory variables are again seen in Table 32 in Appendix B. Compared to the BB experiment above, some further variables are required now. One of them is the difference in inflation rates in the tradables sector between the CEE country and the EU11. Here we consider two cases. The first is that PPP in tradables starts to holds from now on. This implies that the deviation from PPP in tradables term in all the BS-equations is set to zero. The corresponding results are given in Table 17. This reflects the assumption that due to EU enlargement tradables prices should move towards PPP across the enlarged EU. The second assumption, with the corresponding results given in Table 33 in Appendix B is to base the scenario on the historical averages also for the tradables price differences. In the equations displayed in Tables 9 and 10 also some demand variables are present. These are the growth rates of real per capita GDP and of real total consumption.

For these two variables the scenario values are given by the mean prediction for real per capita GDP growth derived in Wagner and Hlouskova (2004). Since population growth is currently close to zero in the CEECs, under the assumption of balanced growth the real per capita GDP growth rate is approximately equal the growth rate of real total consumption. All the listed assumptions suffice for computing the GDP deflator inflation rates for the CEEC8.

A couple of clear observations can be made in Table 17. First of all, the inflation rates are with very few exceptions monotonously declining from the first (1994–2001) to the last (2000–2001) panel for all countries, and subsequently also for the CEEC8 as a group. Comparing the mean of the implied inflation rates (for the CEEC8 group) with the historic values shows that the 'fit' is especially good for the last period. This is noteworthy since the parameters of the equations are estimated for the entire period 1994–2001. Thus, our scenarios are tracking the direction of the disinflation process observed since the mid 1990ies. We focus in the rest of the discussion on the simulation based on the last period values, since they very precisely match the last observations in the sample. This match occurs despite two counterfactual assumptions, namely identical inflation rates in tradables and no nominal exchange rate changes (required for the inflation computations in the equations with Δq or Δq_2 as dependent variables). Note in Table 32, that the averages for the independent variables show no clear trends across the three averaging periods, contrary to e.g. the GDP deflator inflation rates. Thus, loosely speaking, the description of the actual inflation movements by our extended BS-equations becomes more accurate towards the end of the sample. This is, of course, partly at least a level effect that comes from disinflation. However, it does not invalidate the fact that very accurate inflation simulations are obtained, when the fit over the period 2000-2001 is chosen as the evaluation criterion.

The mean inflation projections range from 2.77% for the Slovak Republic to 6.75 % for Poland. The standard deviation varies from about 0.8% for the Slovak Republic to about 3.2% for Lithuania. The standard deviation of the mean simulation for CEEC8 inflation, given by 5.43%, is about 1.2% inflation. Thus, roughly the interval between about 4 to about 6.5% inflation rate is the result of the Balassa-Samuelson inflation projection exercise, for the CEEC8 group.[20]

Note, that (for all three periods) the means over the equations with Δp_2 or Δq_2 as depen-

[20] When the inflation simulations are computed with tradables price differentials set at the historical values, then mean inflation simulation for the CEEC8 is given by 7.92%, with a standard deviation of 0.93% inflation. See Table 33 in Appendix B for details.

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN	CEEC8
1994–2001									
Min	2.89	3.20	3.52	2.71	3.33	3.99	2.78	2.52	3.83
Max	6.63	8.73	7.66	10.31	10.98	10.13	5.21	6.67	7.84
Mean	4.60	6.00	5.44	6.64	6.90	7.16	3.84	4.49	5.99
Std. Dev.	0.92	1.67	1.58	2.36	2.59	2.05	0.80	1.34	1.30
Mean p, q	4.36	4.89	4.24	6.47	6.47	5.92	3.52	3.51	5.05
Mean p_2, q_2	4.87	7.27	6.81	6.83	7.39	8.58	4.20	5.61	7.06
1996–2001									
Min	2.07	3.41	3.55	2.69	1.21	3.92	2.45	2.43	3.68
Max	5.73	6.84	7.06	10.16	9.50	9.60	4.63	6.00	7.23
Mean	4.06	5.12	5.25	6.01	5.74	6.98	3.37	4.16	5.69
Std. Dev.	0.94	1.12	1.31	2.49	2.66	1.87	0.73	1.19	1.09
Mean p, q	4.07	4.76	4.32	6.52	6.38	6.05	3.29	3.34	5.04
Mean p_2, q_2	4.13	5.53	6.32	5.43	5.01	8.05	3.46	5.10	6.44
2000–2001									
Min	2.14	1.52	3.36	1.48	-1.70	3.08	1.64	1.15	3.10
Max	5.73	6.84	7.06	10.58	10.13	9.60	4.63	6.00	7.23
Mean	4.10	3.76	4.96	5.25	4.65	6.75	2.77	3.81	5.43
Std. Dev.	1.02	1.05	1.21	2.54	3.23	2.05	0.81	1.59	1.17
Mean p, q	4.08	3.99	4.30	5.94	6.35	5.55	2.94	2.69	4.73
Mean p_2, q_2	4.11	3.50	5.72	4.48	2.70	8.13	2.59	5.10	6.24

Table 17: Balassa-Samuelson inflation simulations under the assumption Δp^* equals 2% and with the inflation differentials in tradables set to zero. The values for the other variables are at the average values for the periods specified, except for real per capita GDP and real total consumption growth, which are taken from Wagner and Hlouskova (2004).
The three panels correspond to the periods over which the average values for the explanatory variables (except for per capita GDP and total consumption) are taken.
Min, *Max*, *Mean* and *Std.Dev.* denote the minimum, maximum, mean and standard deviation of the implied inflation rates for all 15 equations. *Mean* p, q and *Mean* p_2, q_2 denote the mean over the corresponding sub-groups of equations only.

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN
1994-2001	-0.04	0.12	10.86	-4.36	-4.38	6.85	2.30	6.22
1996-2001	-0.30	0.71	7.40	-3.47	-6.31	2.45	1.80	5.70
2000-2001	-2.20	0.00	0.75	-5.53	-8.71	-7.04	-0.94	5.71

Table 18: Nominal exchange rate changes (in % per annum) of CEEC currencies against the Euro. The exchange rates are defined as units of local currency per Euro. Averages are computed over the periods indicated in the first column.

dent variables are higher, by about 1.5 to 2%, than the means over the equations with Δp or Δq as dependent variables.

The simulations have been based on the assumption of an unchanged nominal exchange rate. Table 18 shows that with the exception of Estonia over the period 2000–2001 no country has experienced a constant exchange rate vis-a-vis the Euro.[21] Furthermore the data support a link between nominal appreciation and inflation reduction over the sample period. It thus might be a valuable exercise to refine the inflation simulations with respect to assumptions concerning nominal exchange rate movements. Here, importantly, the upcoming European Monetary Union membership of the CEE countries represents the anchor for potentially refined simulation exercises. This is, however, beyond the scope of this paper, as it also requires estimates of the pass through of exchange rate changes.

All inflation simulation experiments lead to the conclusion that 2% is a too tight inflation objective for the CEE countries. An aggregate inflation around 4 to 5% seems to be more appropriate over the medium run. Also the inflation differentials across countries are expected to remain substantial. These two facts may present challenges for monetary policy in the enlarged European Monetary Union.

8 Summary and Conclusions

In this paper we offer a detailed econometric analysis of the Baumol-Bowen and Balassa-Samuelson effects for eight CEE countries. Our results are based on a variety of specifications derived from the theoretical model presented in Section 2. We estimate specifications with narrowly (p_2, q_2) and broadly (p^{GDP}, q) defined dependent variables and also employ various BS variables. The narrow specifications result in general in slightly better fit, which is con-

[21]Note again that the exchange rate movements for the Baltic countries stem from the currency board arrangements in place in Estonia, with respect to the SDR, and Latvia, with respect to the Euro since 1999.

sistent with the fact that they are more closely related to the underlying model. We test for the validity of several key assumption of the BB and BS model. These are homogenous sectoral wages, prevalence of PPP in tradables and irrelevance of demand side factors. All three assumptions are refuted by the data. We thus perform our empirical analysis with extended equations that account for the non-validity of these assumptions. Real per capita GDP is found to be the most important demand side variables, this is consistent with Bergstrand (1991).

Based on extensive bootstrap panel unit root and panel cointegration testing, we find throughout evidence for unit root nonstationarities in the data, but no evidence for cointegration. We resort to bootstrapping in order to overcome, or at least mitigate, the bad small sample performance of panel unit root and panel cointegration tests. From the lack of evidence for cointegration we conclude that other studies that rely upon cointegration may have done so inappropriately. In order to assess the possible biases of such practice, we also specify the full set of equations including incorrectly (nonstationary) 'error correction' terms. Taking the differences in the estimated effects between the corresponding equations without and with error correction terms as a measure of the bias, we find that incorrect application of cointegration techniques results for all countries and equations in an overestimation of the BB and BS effects. For the BB effect this is in general by a factor of about four and for the BS effect the average overestimation is by a factor two to four depending upon the specifications considered.

Evidence for the BB and BS effect is found. However, the effects are found to be *small*, about half a percent per annum on average. With the more theory driven measures for the BS variables, the BS effect is even negative for most or all countries. The small magnitude of the effects does not explain the large inflation differentials observed between the CEE countries and the EU11. This is perfectly consistent with the observation that several key assumptions of the standard model are not supported by the data. We therefore base our inflation simulations on the well specified extended equations. Thus, our simulations incorporate the extensions required by the data. We specify several scenarios to obtain inflation simulations based on the BB equation and all the BS equations. The inflation simulations rest on the following assumptions. The independent variables are set at their historical average values (computed over several periods), except for real per capita GDP growth, which is taken from Wagner and Hlouskova (2004) and the deviation from PPP for tradables, which is set to zero for the BS simulations presented in the main text. The inflation rate in the EU11 is set to 2%. The results from the BS inflation simulations can be summarized as follows. An inflation objective of 2% seems to be too low for the CEECs over the medium-run. This finding is also supported by the BB inflation simulations, where an inflation objective of 2% for the CEEC8 leads to deflation in several countries and also to substantial deflation in tradables prices.

The mean inflation predictions range from 2.77% for the Slovak Republic to 6.75% for Poland. The mean inflation prediction for the CEEC8 aggregate inflation is 5.43%. These findings imply that common monetary policy in the enlarged European Monetary Union to come might have to be adjusted to allow for higher and more versatile inflation rates across the CEE countries. Note finally that the results of this paper (the specified equations) can be used to derive additional inflation simulations based on more detailed scenario assumptions. These refinements could be with respect to the nominal exchange rates (set constant in our simulations) or with respect to inflation in agriculture and the public sector (set at historical averages in our simulations). This is, however, beyond the scope of this paper.

References

Alberola, A. and T. Tyrväinen (1998). Is there Scope for Inflation Differentials in EMU? An Empirical Investigation of the Balassa-Samuelson Model in EMU Countries. Bank of Finland Discussion Paper 15/98.

Arellano, M. (2003). *Panel Data Econometrics*, Oxford, Oxford University Press.

Asea, P.K. and E.G. Mendoza (1994). The Balassa-Samuelson Model: A General Equilibirum Appraisal. *Review of International Economics* **2**, 244–267.

Balassa, B. (1964). The Purchasing Power Parity Doctrine: A Reappraisal. *Journal of Political Economy* **72**, 584–596.

Baumol, W. and W. Bowen (1966). *Performing Arts: The Economic Dilemma*, New York, 20th Century Fund.

Bergstrand, J.H. (1991). Structural Determinants of Real Exchange Rates and National Price Levels: Some Empirical Evidence. *American Economic Review* **81**, 325–334.

Breitung, J. (2000). The Local Power of some Unit Root Tests for Panel Data, 161–177. In Baltagi, B.H. (Ed.) *Nonstationary Panels, Panel Cointegration, and Dynamic Panels*, Elsevier, Amsterdam.

Canzoneri, M.B., R.E. Cumby and B. Diba (1999). Relative Labor Productivity and the Real Exchange Rate in the Long Run: Evidence for a Panel of OECD Countries. *Journal of International Economics* **47**, 245–266.

Chang, Y. (2000). Bootstrap Unit Root Tests in Panels with Cross-Sectional Dependency. forthcoming in *Journal of Econometrics*.

Chang, Y. (2004). Nonlinear IV Unit Root Tests in Panels with Cross-Sectional Dependency. *Journal of Econometrics* **110**, 261–292.

Chang, Y. and W. Song (2002). Panel Unit Root Tests in the Presence of Cross-Sectional Dependency and Heterogeneity. Rice University Working Paper N0. 2002-06.

Chiang, M-H. and C. Kao (2002). Nonstatonary Panel Time Series Using NPT: A User Guide. Center for Policy Research, Syracuse University.

Coricelli, F. and B. Jazbec (2004a). Exchange Rate Arrangements in the Accession to the EMU. *Comparative Economic Studies* **46**, 4–22.

Coricelli, F. and B. Jazbec (2004b). Real Exchange Rate Dynamics in Transition Economies. *Structural Change and Economic Dynamics* **15**, 83–100.

Davidson, R. and J.G. MacKinnon (1993). *Estimation and Inference in Econometrics*, New York: Oxford University Press.

Egert, B. (2002). Investigating the Balassa-Samuelson Hypothesis in the Transition. Do We Understand what We See? A Panel Study. *Economics of Transition* **10**, 273–309.

Egert, B., Drine, I., Lommatzsch., K. and Ch. Rault (2002). The Balassa-Samuelson Effect in Central and Eastern Europe: Myth or Reality. William Davidson Institute Working Paper N0. 483.

Fischer, Ch. (2002). Real Currency Appreciation in Accession Countries: Balassa-Samuelson and Investment Demand. Deutsche Bundesbank Discussion Paper No. 19/02.

Ghironi, F. and M. Melitz (2003). International Trade and Macroeconomic Dynamics with Heterogeneous Firms. Mimeo.

Grafe, C. and Ch. Wyplosz (1999). The Real Exchange Rate in Transition Economies. In Blejer, M. and M. Skreb (Eds.), *Balance of Payments, Exchange Rates and Competitiveness in Transition Economies*. Kluwer Academic Publishers.

Gutierrez, L. (2003). On the Power of Panel Cointegration Tests: A Monta Carlo Comparison. *Economics Letters* **80**, 105–111.

Halpern, L. and Ch. Wyplosz (2002). Economic Transformation and Real Exchange Rates in the 2000s: The Balassa-Samuelson Connection. *UNO Economic Surveys of Europe* **1**, 227-239.

Harberger, A.C. (2004). The Real Exchange Rate: Issues of Concept and Measurement. Mimeo.

Harris, R.D.F. and E. Tzavalis (1999). Inference for Unit Roots in Dynamic Panels Where the Time Dimension is Fixed. *Journal of Econometrics* **90**, 1–44.

Hlouskova, J. and M. Wagner (2004a). The Performance of Panel Unit Root Tests: A Simulation Study. Mimeo.

Hlouskova, J. and M. Wagner (2004b). The Performance of Panel Cointegration Methods: A Simulation Study. Mimeo.

Im, K.S., M.H. Pesaran and Y. Shin (1997). Testing for Unit Roots in Heterogeneous Panels. Mimeo.

Im, K.S., M.H. Pesaran and Y. Shin (2003). Testing for Unit Roots in Heterogeneous Panels. *Journal of Econometrics* **115**, 53–74.

Kao, C. (1999). Spurious Regression and Residual Based Tests for Cointegration in Panel Data. *Journal of Econometrics* **90**, 1–44.

Kao, C. and M.-H. Chiang (2000). On the Estimation and Inference of a Cointegrated Regression in Panel Data. In Baltagi, B.H. (Ed.) *Nonstationary Panels, Panel Cointegration, and Dynamic Panels*, Elsevier, Amsterdam.

Kovacs, M.A. (Ed.) (2002). On the Estimated Size of the Balassa-Samuelson Effect in Five Central and Eastern European Countries. National Bank of Hungary Working Paper 2002/5.

Levin, A., C.F. Lin and C-S.J. Chu (2002). Unit Root Tests in Panel Data: Asymptotic and Finite Sample Properties. *Journal of Econometrics* **108**, 1–22.

Maddala, G.S. and S. Wu (1999). A Comparative Study of Unit Root Tests with Panel Data and a Simple New Test. *Oxford Bulletin of Economics and Statistics* **61**, 631–652.

Mark, N.C. and D. Sul (2001). Cointegration Vector Estimation by Panel Dynamic OLS and Long-Run Money Demand. Mimeo.

Mihaljek, D. and M. Klau (2004). The Balassa-Samuelson Effect in Central Europe: A Disaggregated Analysis. *Comparative Economic Studies* **46**, 63–94.

Paparoditis, E. and D. Politis (2002). Bootstrapping Unit Root Tests for Autoregressive Time Series. Forthcoming in *Journal of the American Statistical Association*.

Paparoditis, E. and D. Politis (2003). Residual-based Block Bootstrap for Unit Rot Testing. *Econometrica* **71**, 813–855.

Pedroni, P. (1999). Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors. *Oxford Bulletin of Economics and Statistics* **61**, 653–670.

Pedroni, P. (2000). Fully Modified OLS for Heterogeneous Cointegrated Panels. In Baltagi, B.H. (Ed.) *Nonstationary Panels, Panel Cointegration, and Dynamic Panels*, Elsevier, Amsterdam.

Pedroni, P. (2004). Panel Cointegration. Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis. Forthcoming in *Econometric Theory*.

Phillips, P.C.B. and B.E. Hansen (1990). Statistical Inference in Instrumental Variables Regression with I(1) Processes. *Review of Economic Studies* **57**, 99–125.

Phillips, P.C.B. and H.R. Moon (1999). Linear Regression Limit Theory for Nonstationary Panel Data. *Econometrica* **67**, 1057–1111.

Phillips, P.C.B. and S. Ouliaris (1990). Asymptotic Properties of Residual Based Tests for Cointegration. *Econometrica* **58**, 165–193.

Saikkonen, P. (1991). Asymptotically Efficient Estimation of Cointegrating Regressions. *Econometric Theory* **7**, 1–21.

Samuelson, P. (1964). Theoretical Notes on Trade Problems. *The Review of Economics and Statistics* **46**, 145–154.

Sinn, H.-W. and M. Reutter (2001). The Minimum Inflation Rate for Euroland. NBER Working Paper Series No. 8085.

Wagner, M. and S. Doytchinov (2004). The Balassa-Samuelson Effect in the EU11. In preparation.

Wagner, M. and J. Hlouskova (2004). CEEC Growth Projections: Certainly Necessary and Necessarily Uncertain. Conditionally accepted for publication in *Economics of Transition*.

Appendix A: Data Description

Symbol	Country
CZE	Czech Republic
EST	Estonia
HUN	Hungary
LVA	Latvia
LTU	Lithuania
POL	Poland
SVK	Slovak Republic
SVN	Slovenia
EU11 countries	
AUT	Austria
BEL	Belgium
DNK	Denmark
FIN	Finland
FRA	France
GER	Germany
GBR	Great Britain
ITA	Italy
NLD	The Netherlands
ESP	Spain
SWE	Sweden

Table 19: List of countries used in this study and abbreviations.

NACE code	NACE category	Sector
A	Agriculture	AGR
B	Fishing	AGR
C	Mining and quarrying	T
D	Manufacturing	T
E	Electricity, gas and water supply	T
F	Construction	N
G	Wholesale and retail trade	N
H	Hotels and restaurants	N
I	Transport, storage and communication	N
J	Financial intermediation	N
K	Real estate and business activities	N
L	Public administration and defence	PUB
M	Education	PUB
N	Health and social work	PUB
O	Other communal, social and indiv. services	PUB
P	Private households with employed persons	PUB

Table 20: Aggregation of NACE categories to the 4 sectors agriculture (AGR), tradables (T), non-tradables (N) and public sector (PUB) as defined in this study.

Symbol	Variable
GDP	Gross domestic product,
	1995 prices, local currency
GDPPC	GDP per capita, constant 1999 US$ (EKS PPP)
HHC	Final consumption of households
	1995 prices, local currency
GOV	Government final consumption
	1995 prices, local currency
NPH	Final consumption of non-profit organizations
	1995 prices, local currency
GFCF	Gross fixed capital formation
	(including changes in inventories)
	1995 prices, local currency
FDI	Foreign direct investment
	net inflows, % of GDP
EXP	Exports of goods and services
	1995 prices, local currency
IMP	Imports of goods and services
	1995 prices, local currency
Y^X	Gross Value Added, 1995 producer prices, local currency
P^Z	Deflators, 1995 = 100, based on local currencies
L^X	Employment, annual average
U	Registered unemployment, total
W^X	Annual gross wages per employee, current prices, local currency
LC^X	Annual labor costs per employee, current prices, local currency
	Labor cost is the sum of gross wages and social security contributions
E	Nominal exchange rate, Local currency/EURO(ECU)

Table 21: List of variables. The super-script X indicates the sector $\{T, N, AGR, PUB\}$, and the super-script Z for the price deflator indicates a value in the set $\{GDP, HHC, GVC, NPH, GFCF, EXP, IMP, T, N, AGR, PUB\}$. No super-script for these variables indicates the economy-wide variables.

Variable	Country	Source
GDP, HHC, GOV, NPH, GFCF, EXP, IMP, P^{GDP}, P^{HHC}, P^{GVC}, P^{NPH}, P^{GFCF}, P^{EXP}, P^{IMP}	EU11 without FIN and FRA FIN, FRA CZE, HUN, EST, LVA, POL, SVK, SVN LTU	WIFO EUROSTAT UNECE WDI (1993-1994), UNECE (1994-2001)
GDPPC	CZE, HUN, EST, LTU, LVA POL, SVK, SVN	Groningen Growth and Development Center at the University of Groningen
FDI	CZE, HUN, EST, LTU, LVA POL, SVK, SVN	World Development Indicators
Y^T, Y^N, Y^{AGR}, Y^{PUB}, Y, P^T, P^N, P^{AGR}, P^{PUB}, P	EU11 CZE, EST, LVA, POL, SVK, SVN HUN, LTU	EUROSTAT EUROSTAT UNECE (1993-1995) EUROSTAT (1996-2001)
L^T, L^N, L^{AGR}, L^{PUB}, L	EU11 EST, LVA, LTU, POL, SVK CZE HUN SVN	EUROSTAT EUROSTAT UNECE UNECE (1993-1995) EUROSTAT (1996-2001) National Statistical Office
U	EU11 CZE, HUN, SVK EST, LVA, LTU, POL SVN	WIFO EUROSTAT UNECE National Statistical Office
W^T, W^N, W^{AGR}, W^{PUB}, W	EU11 CZE EST, HUN, LVA, LTU, POL, SVN SVK	EUROSTAT EUROSTAT (1993-2000) UNECE (2001) EUROSTAT National Statistical Office
LC^T, LC^N, LC^{AGR}, LC^{PUB}, LC	EU11 EST, HUN, LVA, LTU, POL, SVK CZE SVN	EUROSTAT EUROSTAT EUROSTAT (1993-2000) UNECE (2001) National Statistical Office
E	CEE	EUROSTAT

Table 22: Description of data sources. UNECE denotes United Nations Economic Commission for Europe, WDI denotes the World Development Indicators and WIFO denotes the Austrian Institute for Economic Research.

Symbol	Definition
Prices	
p^{GDP}	$ln(P^{GDP})$
δ	$\frac{Y^N}{Y^T+Y^N}$
P^{T+N}	$(1-\delta)P^T + \delta P^N$
p^{T+N}	$ln(P^{T+N})$
p^{rel}	$ln(100P^N/P^T)$
p^T	$ln(P^T)$
p^N	$ln(P^N)$
p^{AGR}	$ln(P^{AGR})$
p^{PUB}	$ln(P^{PUB})$
Labor shares in tradables and non-tradables sectors	
α^T	$LC^T L^T/Y^T$
α^N	$LC^N L^N/Y^N$
Labor productivities	
A^T	Y^T/L^T
A^N	Y^N/L^N
Ai^T	$100A^T/A^T_{1995}$
Ai^N	$100A^N/A^N_{1995}$
a^{rel}	$ln(100Ai^T/Ai^N)$
a^T	$ln(Ai^T)$
a^N	$ln(Ai^N)$
a^{rel}_m	$ln(100(Ai^T)^{\alpha^N/\alpha^T}/Ai^N)$
Wages and labor costs	
Wi^T	$100W^T/W^T_{1995}$
Wi^N	$100W^N/W^N_{1995}$
w^{rel}	$ln(100Wi^N/Wi^T)$
w^T	$ln(Wi^T)$
w^N	$ln(Wi^N)$
w^{rel}_{BS}	$w^{rel} - w^{rel*} + ln(100)$
LCr^T	$100LC^T/P^T$
LCr^N	$100LC^N/P^N$
lcr^T	$ln(100LCr^T/LCr^T_{1995})$
lcr^N	$ln(100LCr^N/LCr^N_{1995})$
Real exchange rates	
Q	$EP^{GDP}_{EU11}/P^{GDP}_{CEE}$
q	$ln(100Q/Q_{1995})$
Q_2	$EP^{T+N}_{EU11}/P^{T+N}_{CEE}$
q_2	$ln(100Q_2/Q_{2,1995})$
Q^T	EP^T_{EU11}/P^T_{CEE}
q^T	$ln(100Q^T/Q^T_{1995})$
Total consumption	
TC	$HHC + GOV + NPH$

Table 23: Detailed description of variable transformation employed in the empirical analysis.

Gross domestic product, 1995, EURO	
$GDP^{*} = \sum_{i \in C} GDP_i / E_{i,1995}$	
Gross value added, 1995 producer prices, EURO	
$Y^{T*} = \sum_{i \in C} Y_i^T / E_{i,1995}$	$Y^{N*} = \sum_{i \in C} Y_i^N / E_{i,1995}$
Employment	
$L^{T*} = \sum_{i \in C} L_i^T$	$L^{N*} = \sum_{i \in C} L_i^N$
GDP weights	
$c_i = \frac{GDP_i / E_{i,1995}}{GDP^{*}}, \, i \in C$	
Sectoral value added weights	
$c_i^T = \frac{Y_i^T / E_{i,1995}}{Y^{T*}}, \, i \in C$	$c_i^N = \frac{Y_i^N / E_{i,1995}}{Y^{N*}}, \, i \in C$
Deflators, 1995=100	
$P^{GDP*} = \sum_{i \in C} c_i P_i^{GDP} E_{i,1995} / E_i$	
$P^{T*} = \sum_{i \in C} c_i^T P_i^T E_{i,1995} / E_i$	$P^{N*} = \sum_{i \in C} c_i^N P_i^N E_{i,1995} / E_i$
Labor productivities	
$A^{T*} = Y^{T*} / L^{T*}$	$A^{N*} = Y^N_{EU11} / L^{N*}$
Annual gross wages per employee, current prices, Euro	
$W^{T*} = \frac{\sum_{i \in C} (W_i^T / E_i) L_i^T}{L^{T*}}$	$W^{N*} = \frac{\sum_{i \in C} (W_i^N / E_i) L_i^N}{L^{N*}}$
Annual labor costs per employee, current prices, Euro	
$LC^{T*} = \frac{\sum_{i \in C} (LC_i^T / E_i) L_i^T}{L^{T*}}$	$LC^{N*} = \frac{\sum_{i \in C} (LC_i^N / E_i) L_i^N}{L^{N*}}$

Table 24: Details of construction of the variables for the EU11 aggregate. C here denotes the index set comprising eleven countries.

Appendix B: Additional Empirical Results

	LL	UB	IPS	HT	$IPS-LM$	MW
q	**-2.593*** (-6.816)	0.906 (-0.138)	**-2.348** (-1.676)	**-0.048** (0.718)	**1.048** (1.014)	**62.572** (39.059)
q_2	**-2.462*** (-5.523)	1.131 (-0.017)	-0.905 (-1.136)	**0.244** (0.948)	0.272 (0.803)	**36.435** (34.020)
q^T	-1.083 (-5.089)	0.294 (-0.462)	**-1.660*** (-2.010)	**-0.254** (0.460)	0.916 (1.333)	**39.250** (39.084)
p^{rel}	3.235 (-6.275)	1.538 (-0.272)	0.212 (-1.329)	**0.021** (0.233)	-0.465 (0.723)	19.091 (42.112)
$p^T - p^{T*}$	**-4.001*** (-4.333)	0.509 (-0.181)	**-4.902** (-2.322)	0.303 (0.893)	**2.138** (1.478)	**121.282** (41.179)
$p^{GDP} - p^{GDP*}$	1.474 (-4.638)	**-0.983** (-0.223)	**-5.177** (-3.306)	**0.161** (0.765)	**2.382*** (2.667)	**121.451** (42.191)
$p^{T+N} - p^{(T+N)*}$	**-2.606*** (-3.665)	-0.151 (-0.165)	**-4.334** (-2.762)	**0.475** (1.070)	**2.189*** (2.197)	**94.208** (42.901)
w^T	**-3.371*** (-5.886)	1.521 (-0.265)	**-2.022** (-1.604)	**0.891** (1.455)	1.481 (1.593)	**39.818** (31.104)
w^N	**-9.505** (-4.286)	1.927 (-0.200)	**-1.652** (-1.024)	**1.386** (1.791)	**1.039** (0.840)	**52.027** (29.534)
w^{rel}	-1.109 (-5.095)	-0.807 (-0.955)	-0.937 (-1.917)	**-4.210** (-1.845)	0.484 (1.050)	**33.806*** (44.750)
w^{rel}_{BS}	**-1.656*** (-5.108)	-1.133 (-1.138)	-0.726 (-2.219)	**-4.236** (-1.706)	0.353 (1.028)	**32.950*** (45.765)
a^T	-0.925 (-4.859)	0.820 (-0.101)	1.471 (0.048)	1.823 (1.784)	-1.724 (0.348)	7.510 (22.786)
a^N	**-2.066*** (-3.603)	1.573 (-0.303)	3.362 (-0.079)	2.407 (1.399)	-2.373 (0.466)	2.756 (30.115)
a^{rel}	**-2.658*** (-4.408)	0.084 (-0.484)	-0.488 (-1.072)	**0.039** (0.226)	0.264 (0.786)	18.226 (30.339)
$a^{rel} - a^{rel*}$	**-2.658*** (-4.408)	-0.287 (-0.639)	**-2.043*** (-4.214)	-0.529 (-0.658)	**1.247** (0.916)	**77.414** (33.713)
BS	4.468 (-9.387)	-1.227 (-1.577)	1.154 (-3.564)	-1.619 (-2.075)	-1.564 (1.491)	6.869 (61.403)
$BSE1$	1.210 (-4.351)	0.579 (-0.734)	1.948 (-2.134)	0.745 (-0.440)	-2.025 (1.188)	5.527 (41.724)
$BSE2$	1.106 (-9.042)	0.492 (-1.049)	2.906 (-2.219)	0.363 (-1.218)	-1.793 (1.133)	5.581 (46.907)

Table 25: Results of panel unit root tests including fixed effects. In parentheses the 5% critical values obtained by the *non-parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by -1.645 for the first 4 tests, by 1.645 for IPS-LM and by 26.296 for MW.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in the autoregression based tests is equal to one and the autoregressive lag lengths the non-parametric bootstrap are equal to one.

	LL	UB	IPS	HT	$IPS-LM$	MW
q	0.650 (-14.391)	-0.544 (-1.074)	-1.111 (-5.739)	1.122 (-3.280)	0.180 (1.263)	**41.843*** (89.986)
q_2	1.945 (-14.848)	-0.436 (-0.983)	**-3.249*** (-5.090)	0.752 (-3.065)	0.420 (1.348)	**109.131** (74.472)
q^T	-0.862 (-16.963)	-0.852 (-1.013)	**-1.818*** (-4.879)	0.088 (-3.107)	0.684 (1.304)	**47.933*** (66.648)
p_{rel}	0.334 (-16.469)	-0.664 (-0.817)	1.328 (-4.601)	-1.455 (-2.784)	-1.834 (1.211)	13.641 (79.761)
$p^T - p^{T*}$	**-2.481*** (-13.756)	0.114 (-1.376)	-1.035 (-6.087)	2.429 (-3.240)	0.488 (1.549)	**30.208*** (93.244)
$p^{GDP} - p^{GDP*}$	5.423 (-19.295)	**-1.671*** (-1.947)	**-7.038** (-6.961)	2.871 (-4.480)	0.890 (2.297)	**53.705*** (95.509)
$p^{T+N} - p^{(T+N)*}$	0.784 (-14.449)	0.071 (-1.718)	**-4.085*** (-6.616)	2.992 (-4.124)	0.746 (2.060)	**119.656** (86.616)
w^T	0.812 (-14.840)	0.068 (-1.454)	0.411 (-5.378)	2.453 (-3.489)	-0.954 (1.762)	20.508 (79.480)
w^N	-1.192 (-18.359)	0.864 (-1.347)	0.769 (-5.376)	2.600 (-3.027)	-1.033 (1.427)	13.311 (82.905)
w^{rel}	**-6.616*** (-13.860)	0.627 (-1.106)	**-4.195*** (-4.642)	**-2.462*** (-2.670)	-0.047 (1.162)	**48.768*** (71.407)
w^{rel}_{BS}	**-5.343*** (-13.885)	0.473 (-1.163)	**-2.267*** (-4.765)	**-2.506*** (-2.639)	-0.328 (1.136)	**94.633** (66.774)
a^T	**-13.737** (-13.023)	-0.661 (-1.068)	**-4.166*** (-4.428)	-0.425 (-2.561)	1.083 (1.290)	**117.895** (62.003)
a^N	0.372 (-14.055)	-0.192 (-0.824)	1.030 (-4.271)	-0.807 (-2.553)	-1.187 (1.307)	9.801 (68.624)
a^{rel}	**-11.985*** (-16.324)	-0.829 (-1.188)	-1.280 (-4.040)	0.220 (-2.597)	-0.009 (1.247)	**52.175*** (60.893)
$a^{rel} - a^{rel*}$	**-11.985*** (-16.324)	0.299 (-1.081)	**-2.043*** (-4.214)	0.152 (-2.612)	0.355 (1.259)	**68.918** (61.989)
BS	**-8.828*** (-24.857)	0.357 (-0.813)	**-5.126*** (-6.034)	**-2.081*** (-2.679)	0.617 (1.315)	**157.148** (73.746)
$BSE1$	3.468 (-14.689)	-0.180 (-0.859)	0.291 (-4.738)	-0.194 (-2.441)	-1.089 (1.249)	**30.413*** (74.875)
$BSE2$	-0.120 (-18.500)	0.472 (-0.785)	1.158 (-4.803)	0.073 (-2.678)	-0.934 (1.394)	8.957 (79.758)

Table 26: Results of panel unit root tests including fixed effects and time trends. In parentheses the 5% critical values obtained by the *non-parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by -1.645 for the first 4 tests, by 1.645 for IPS-LM and by 26.296 for MW.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in the autoregression based tests are equal to one and the autoregressive lag lengths in non-parametric bootstrap is equal to one.

	PP_σ	PP_ρ	PP_t	PP_{df}	PG_ρ	PG_t	PG_{df}
$Wages$	**3.228** (2.770)	-1.532 (-1.931)	**-3.169*** (-4.475)	**-2.619*** (-4.082)	-0.160 (-0.316)	**-3.224*** (-4.547)	**-3.666*** (-5.297)
LC-LPT	1.053 (1.888)	-1.143 (-2.322)	**-2.666*** (-5.652)	**-2.258*** (-5.122)	0.196 (-0.661)	**-2.813*** (-6.078)	**-2.977*** (-6.560)
LC-LPN	1.541 (2.577)	-1.399 (-2.096)	**-4.463*** (-4.973)	**-3.758*** (-4.496)	0.091 (-0.415)	**-5.107** (-5.023)	**-5.093*** (-5.597)
BBE	-0.414 (1.254)	0.369 (-0.489)	-0.836 (-4.357)	-0.430 (-3.842)	1.403 (0.904)	-1.266 (-4.599)	**-1.677*** (-5.109)
Aq	-1.324 (0.483)	1.271 (-0.680)	-0.008 (-5.115)	0.516 (-4.509)	1.718 (0.793)	-0.981 (-5.322)	-1.524 (-5.718)
Aq_2	-1.610 (0.655)	1.448 (-0.720)	0.489 (-5.243)	0.963 (-4.596)	1.855 (0.786)	-0.764 (-5.340)	-1.274 (-5.756)
Ap	-1.113 (0.705)	0.752 (-0.492)	-0.899 (-4.452)	-0.579 (-4.014)	1.397 (0.910)	**-1.778*** (-4.725)	**-2.406*** (-5.353)
Ap_2	-1.238 (0.976)	0.775 (-0.461)	-0.714 (-4.298)	-0.442 (-3.877)	1.310 (0.896)	**-1.957*** (-4.822)	**-2.588*** (-5.380)
Bq	-1.878 (-0.196)	1.120 (-0.558)	-0.381 (-4.795)	0.296 (-4.313)	1.872 (0.887)	-0.377 (-4.905)	-0.828 (-5.375)
Bq_2	-2.098 (0.058)	1.616 (-0.573)	0.775 (-4.761)	1.370 (-4.241)	1.969 (0.858)	0.083 (-4.968)	-0.322 (-5.515)
Bp	-0.097 (0.643)	-0.340 (-0.606)	**-3.559*** (-4.896)	**-3.233*** (-4.399)	**0.814** (0.843)	**-3.376*** (-5.021)	**-4.096*** (-5.858)
Bp_2	-0.011 (0.822)	-0.297 (-0.609)	**-2.957*** (-4.888)	**-2.815*** (-4.290)	**0.773** (0.871)	**-2.950*** (-4.889)	**-3.704*** (-5.727)
Cq	-2.840 (0.116)	2.061 (0.496)	-0.682 (-4.552)	-0.382 (-3.921)	2.523 (1.925)	-0.909 (-4.889)	-1.451 (-5.336)
Cq_2	-2.839 (0.034)	2.032 (0.490)	-0.594 (-4.702)	-0.405 (-4.092)	2.680 (1.923)	-0.342 (-4.860)	-0.868 (-5.441)
Cp	-0.666 (0.921)	0.646 (0.315)	**-3.224*** (-5.760)	**-2.989*** (-4.980)	1.908 (1.694)	**-3.673*** (-6.564)	**-4.633*** (-6.966)
Cp_2	-0.747 (0.966)	0.682 (0.307)	**-3.124*** (-5.755)	**-2.952*** (-4.971)	1.910 (1.696)	**-3.576*** (-6.513)	**-4.458*** (-6.974)
Dq	-1.418 (-0.098)	1.627 (0.669)	-0.754 (-3.999)	0.406 (-3.466)	2.336 (2.076)	-1.063 (-3.892)	**-1.723*** (-4.582)
Dq_2	-1.329 (0.021)	1.752 (0.677)	-0.250 (-3.872)	0.790 (-3.336)	2.354 (2.073)	-1.252 (-4.049)	**-1.891*** (-4.709)
Dp	-1.186 (0.341)	1.711 (0.593)	-0.324 (-4.221)	0.826 (-3.707)	2.172 (1.925)	**-2.260*** (-4.699)	**-3.042*** (-5.512)
Dp_2	-1.589 (0.400)	1.899 (0.620)	0.163 (-4.031)	1.400 (-3.567)	2.134 (1.952)	**-2.786*** (-4.556)	**-3.480*** (-5.359)
Eq	-0.354 (0.496)	1.759 (1.601)	**-1.811*** (-3.934)	-1.265 (-3.075)	3.038 (2.928)	**-2.227*** (-3.735)	**-2.430*** (-4.797)
Eq_2	-0.534 (0.471)	1.574 (1.528)	**-2.537*** (-3.853)	**-2.220*** (-3.460)	2.894 (2.849)	**-2.453*** (-4.285)	**-2.957*** (-5.394)
Ep	0.010 (0.475)	**1.347** (1.421)	**-4.172*** (-4.660)	**-3.782*** (-4.096)	**2.639** (2.754)	**-4.256*** (-5.323)	**-4.740*** (-5.956)
Ep_2	-0.723 (0.477)	**1.207** (1.447)	**-4.660** (-4.498)	**-4.646** (-3.921)	**2.574** (2.774)	**-4.842*** (-5.201)	**-5.977** (-5.859)

Table 27: Results of Pedronis panel cointegration tests including fixed effects. In parentheses the 5% critical values obtained by the *parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by 1.645 for the first test and by -1.645 for the other 6 tests.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both the autoregression based tests and the parametric bootstrap are equal to one. The window-length of the Bartlett kernel for the non-parametric tests is also equal to one.

	PP_σ	PP_ρ	PP_t	PP_{df}	PG_ρ	PG_t	PG_{df}
$Wages$	**0.922** (0.714)	**-0.551** (-0.173)	**-6.123** (-4.939)	**-5.681** (-4.555)	**0.762** (1.090)	**-6.317** (-4.844)	**-6.628** (-5.828)
LC-LPT	-0.313 (0.620)	0.636 (-0.121)	**-2.287*** (-4.742)	**-1.768*** (-4.259)	1.711 (1.145)	**-1.927*** (-4.588)	**-2.482*** (-5.496)
LC-LPN	-0.405 (0.608)	0.523 (-0.109)	**-2.761*** (-4.755)	**-2.452*** (-4.243)	1.727 (1.154)	**-2.456*** (-4.532)	**-4.170*** (-5.482)
BBE	-0.540 (0.017)	1.120 (0.899)	**-3.401*** (-4.572)	**-3.120*** (-4.018)	2.258 (2.057)	**-3.386*** (-4.586)	**-3.858*** (-5.392)
Aq	-0.480 (-0.000)	1.165 (0.865)	**-2.794*** (-4.733)	**-2.264*** (-4.164)	2.276 (2.041)	**-2.365*** (-4.951)	**-3.045*** (-5.789)
Aq_2	-0.461 (-0.006)	1.399 (0.873)	-1.547 (-4.742)	-1.075 (-4.266)	2.504 (2.038)	-1.187 (-4.989)	**-1.759*** (-5.776)
Ap	-0.536 (0.047)	1.321 (0.843)	**-2.773*** (-4.874)	**-1.983*** (-4.311)	2.359 (2.004)	**-2.857*** (-5.001)	**-3.382*** (-5.841)
Ap_2	-0.476 (0.057)	1.314 (0.824)	**-2.037*** (-4.923)	**-1.768*** (-4.350)	2.415 (1.987)	-1.349 (-5.035)	**-2.135*** (-5.978)
Bq	-0.520 (0.027)	1.341 (0.844)	**-2.187*** (-4.848)	**-1.793*** (-4.339)	2.467 (2.012)	-1.580 (-4.903)	**-2.133*** (-5.735)
Bq_2	-0.464 (0.044)	1.366 (0.838)	-1.624 (-4.848)	-1.462 (-4.311)	2.521 (1.999)	-0.540 (-4.999)	-1.196 (-5.796)
Bp	-0.067 (0.069)	0.980 (0.820)	**-3.674*** (-5.011)	**-3.371*** (-4.476)	2.144 (1.985)	**-2.899*** (-5.177)	**-3.571*** (-5.973)
Bp_2	-0.181 (0.074)	**0.739** (0.806)	**-4.226*** (-5.018)	**-4.124*** (-4.463)	**1.921** (1.973)	**-3.002*** (-5.168)	**-3.892*** (-6.070)
Cq	-1.441 (-0.583)	2.183 (1.744)	**-1.977*** (-4.627)	**-1.891*** (-4.119)	3.169 (2.888)	-1.318 (-5.038)	**-2.257*** (-5.905)
Cq_2	-1.508 (-0.581)	2.179 (1.742)	**-2.287*** (-4.573)	-1.318 (-4.573)	3.162 (2.880)	**-2.031*** (-4.955)	**-2.699*** (-5.827)
Cp	-1.201 (-0.483)	2.391 (1.705)	-0.730 (-4.943)	-0.272 (-4.378)	3.411 (2.845)	0.206 (-5.310)	-0.527 (-6.147)
Cp_2	-1.203 (-0.489)	2.385 (1.684)	-0.804 (-5.025)	-0.566 (-4.434)	3.458 (2.823)	0.292 (-5.249)	-0.449 (-6.110)
Dq	-1.123 (-0.550)	2.003 (1.780)	**-2.507*** (-4.234)	**-1.963*** (-3.649)	3.150 (2.921)	-1.583 (-4.471)	**-2.322*** (-5.434)
Dq_2	-1.059 (-0.550)	2.126 (1.769)	-1.558 (-4.317)	-1.429 (-3.770)	3.262 (2.908)	-0.363 (-4.572)	-1.119 (-5.569)
Dp	-0.755 (-0.470)	1.989 (1.641)	**-2.656*** (-5.276)	**-2.272*** (-4.662)	3.119 (2.796)	**-1.841*** (-5.674)	**-2.622*** (-6.433)
Dp_2	-0.787 (-0.435)	1.926 (1.663)	**-2.899*** (-5.134)	**-2.657*** (-4.540)	3.035 (2.798)	**-2.170*** (-5.367)	**-2.957*** (-6.075)
Eq	-1.995 (-0.991)	2.930 (2.482)	-1.005 (-4.355)	-0.346 (-3.700)	4.046 (3.632)	-0.464 (-4.858)	-1.232 (-5.766)
Eq_2	-2.005 (-0.825)	2.769 (2.486)	**-1.825*** (-4.336)	-1.125 (-3.744)	3.893 (3.629)	**-1.795*** (-4.834)	**-2.762*** (-5.835)
Ep	-1.088 (-0.963)	2.685 (2.418)	**-2.548*** (-5.099)	**-1.940*** (-4.376)	3.805 (3.566)	**-2.257*** (-5.644)	**-3.034*** (-6.449)
Ep_2	-1.205 (-0.959)	**2.401** (2.418)	**-4.763*** (-5.105)	**-3.798*** (-4.367)	**3.496** (3.566)	**-5.384*** (-5.607)	**-6.420*** (-6.540)

Table 28: Results of Pedronis panel cointegration tests including fixed effects and time trends. In parentheses the 5% critical values obtained by the *parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by 1.645 for the first test and by -1.645 for the other 6 tests.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both the autoregression based tests and the parametric bootstrap are equal to one. The window-length of the Bartlett kernel for the non-parametric tests is also equal to one.

	K_ρ	K_t	K_{df}
$Wages$	**-2.539*** (-2.589)	**-5.003** (-3.270)	-1.541 (-3.064)
LC-LPT	**-2.222*** (-2.390)	**-4.035** (-3.569)	-0.283 (-2.417)
LC-LPN	**-3.466** (-2.183)	**-5.317** (-2.900)	-1.626 (-2.533)
BBE	**-1.993*** (-2.573)	**-3.288** (-2.989)	0.882 (-2.433)
Aq	-0.960 (-3.630)	-1.601 (-4.143)	-1.478 (-4.320)
Aq_2	-0.525 (-3.405)	-1.061 (-3.967)	-1.458 (-4.619)
Ap	-1.299 (-2.710)	**-2.098*** (-3.461)	1.537 (-2.584)
Ap_2	-0.348 (-2.767)	-0.783 (-3.662)	1.980 (-2.617)
Bq	-1.084 (-3.885)	-1.522 (-4.102)	**-2.192*** (-4.530)
Bq_2	-0.837 (-3.640)	-1.132 (-3.971)	**-2.487*** (-4.559)
Bp	-1.150 (-3.393)	**-1.851*** (-3.755)	1.254 (-3.557)
Bp_2	-0.530 (-3.386)	-1.052 (-3.769)	1.508 (-3.747)
Cq	**-2.230*** (-4.077)	**-2.502*** (-4.380)	**-2.204*** (-4.071)
Cq_2	**-1.678*** (-4.067)	**-1.883*** (-4.251)	**-1.847*** (-4.215)
Cp	**-1.902*** (-3.344)	**-3.136*** (-3.901)	0.535 (-2.918)
Cp_2	-1.142 (-3.346)	**-1.914*** (-4.149)	1.380 (-2.219)
Dq	-0.927 (-3.886)	-1.415 (-4.378)	-0.978 (-3.792)
Dq_2	-0.698 (-3.444)	-1.091 (-4.040)	-1.060 (-3.654)
Dp	-1.171 (-3.124)	**-1.900*** (-3.369)	1.835 (-3.646)
Dp_2	-0.690 (-2.859)	-1.044 (-3.137)	2.167 (-3.423)
Eq	-0.975 (-3.482)	-1.443 (-3.701)	-1.234 (-3.882)
Eq_2	-0.933 (-3.156)	-1.409 (-3.550)	-1.212 (-3.719)
Ep	-1.560 (-3.422)	**-2.648*** (-4.030)	1.186 (-3.918)
Ep_2	**-1.646*** (-3.266)	**-2.766*** (-3.853)	1.240 (-3.618)

Table 29: Results of Kao's panel cointegration tests including fixed effects. In parentheses the 5% critical values obtained by the *parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by -1.645 for all 3 tests.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both K_{df} and the parametric bootstrap are equal to one. The Bartlett kernel with window size one is applied.

	K_ρ	K_t	K_{df}
$Wages$	**-2.539** (-1.981)	**-5.003** (-2.726)	-1.541 (-2.424)
LC-LPT	**-2.222** (-1.619)	**-4.035** (-2.439)	-0.283 (-1.503)
LC-LPN	**-3.466** (-1.677)	**-5.317** (-2.298)	-1.626 (-1.872)
BBE	**-1.993*** (-2.278)	**-3.288** (-2.622)	0.882 (-1.974)
Aq	-0.960 (-1.844)	-1.601 (-2.634)	-1.478 (-2.410)
Aq_2	-0.525 (-1.557)	-1.061 (-2.507)	-1.458 (-2.281)
Ap	-1.299 (-1.594)	**-2.098** (-2.085)	1.537 (-1.255)
Ap_2	-0.348 (-1.479)	-0.783 (-2.091)	1.980 (-1.276)
Bq	-1.084 (-2.057)	-1.512 (-2.469)	**-2.192*** (-2.573)
Bq_2	-0.837 (-1.518)	-1.132 (-2.081)	**-2.487** (-2.243)
Bp	-1.150 (-1.850)	**-1.851*** (-2.268)	1.254 (-1.984)
Bp_2	-0.530 (-1.690)	-1.052 (-2.063)	1.508 (-1.884)
Cq	**-2.230*** (-3.269)	**-2.502*** (-3.502)	**-2.204*** (-2.559)
Cq_2	**-1.678*** (-3.200)	**-1.883*** (-3.502)	**-1.847*** (-2.707)
Cp	**-1.909*** (-2.014)	**-3.136** (-2.236)	0.535 (-1.205)
Cp_2	-1.142 (-2.052)	**-1.914*** (-2.241)	1.380 (-0.799)
Dq	-0.927 (-2.904)	-1.415 (-3.616)	-0.978 (-2.403)
Dq_2	-0.698 (-2.898)	-1.091 (-3.714)	-1.060 (-2.504)
Dp	-1.171 (-2.360)	**-1.900*** (-2.747)	1.835 (-2.445)
Dp_2	-0.690 (-2.658)	-1.044 (-2.979)	2.167 (-2.840)
Eq	-0.975 (-2.035)	-1.443 (-2.032)	-1.234 (-1.476)
Eq_2	-0.933 (-2.029)	-1.409 (-2.122)	-1.212 (-1.267)
Ep	-1.560 (-2.585)	**-2.648*** (-2.764)	1.186 (-1.922)
Ep_2	**-1.646*** (-2.732)	**-2.766*** (-2.856)	1.240 (-1.928)

Table 30: Results of Kao's panel cointegration tests including fixed effects. In parentheses the 5% critical values obtained by the *non-parametric* bootstrap are displayed.
The asymptotic 5% critical value is given by -1.645 for all 3 tests.
Bold indicates rejection of the null hypothesis based on the bootstrap critical values and **bold*** indicates rejection based upon the asymptotic critical values but no rejection according to the bootstrap critical values.
The autoregressive lag lengths in both K_{df} and the non-parametric bootstrap are equal to 1. The Bartlett kernel with window size one is applied.

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN	CEEC8
1991	30.89	84.4	–	94.09	113.05	40.46	29.75	66.75	–
1992	11.65	227.59	18.95	237.58	227.01	30.76	10.47	112.55	42.37
1993	19.07	59.59	19.22	53.97	140.17	26.73	14.29	31.54	27.31
1994	12.58	33.37	18.13	32.45	48.02	31.65	12.81	20.37	24.12
1995	9.73	26.95	24.13	13.96	32.24	24.67	9.47	14.11	20.01
1996	8.43	20.92	19.21	15.37	22.38	17.16	4.30	10.53	14.62
1997	7.73	10.09	16.95	6.98	12.41	13.11	6.50	8.44	11.81
1998	10.1	8.90	11.90	5.60	6.46	11.14	5.07	7.54	10.18
1999	2.91	4.42	8.15	6.89	3.19	6.57	6.23	6.37	6.09
2000	1.05	6.70	9.31	4.53	1.96	6.64	6.21	5.58	5.92
2001	5.15	4.92	8.26	1.73	0.26	4.05	5.24	9.43	5.26
1994–2001	7.21	14.53	14.51	10.94	15.87	14.37	6.98	10.30	12.25
1996–2001	5.90	9.33	12.30	6.85	7.78	9.78	5.59	7.98	8.98
2000–2001	3.10	5.81	8.79	3.13	1.11	5.35	5.73	7.51	5.59

Table 31: GDP deflator based inflation rates for the CEEC8 countries and for the group CEEC8.

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN
Δw^{rel}								
1994-2001	0.13	2.16	-1.97	-2.71	-6.84	1.73	0.28	-1.31
1996-2001	0.19	0.58	0.15	2.57	-3.18	3.68	0.23	-1.34
2000-2001	4.68	-3.35	-3.17	0.01	1.82	9.53	-1.73	-1.46
Δw^{rel}_{BS}								
1994-2001	1.32	2.49	-3.09	-5.29	-6.91	-1.46	-0.14	-1.26
1996-2001	1.39	0.24	-1.47	-0.47	-3.01	-4.18	-0.84	-1.74
2000-2001	3.78	-4.24	-4.06	-0.89	0.92	-10.43	-2.63	-2.36
$\Delta p^T - \Delta p^{T*}$								
1994-2001	5.18	11.28	10.52	10.40	16.41	7.95	5.98	8.22
1996-2001	4.75	6.59	8.76	4.22	10.67	5.07	4.12	5.90
2000-2001	6.11	3.52	5.63	0.07	10.92	1.38	2.47	4.62
Δp^{A+P}								
1994-2001	10.57	16.12	14.36	13.05	17.23	16.34	6.07	10.75
1996-2001	8.41	10.32	12.26	7.52	8.61	13.65	3.67	9.30
2000-2001	8.08	5.24	10.25	5.80	0.14	16.56	2.42	10.98
$\Delta ln(GDPPC)$								
	3.22	3.42	3.29	3.34	3.31	3.21	3.33	3.05

Table 32: Period averages of explanatory variables used in the inflation simulations. The real per capita GDP growth rates are the mean projections from Wagner and Hlouskova (2004).

	CZE	EST	HUN	LVA	LTU	POL	SVK	SVN	CEEC8
1994–2001									
Min	7.51	13.04	9.71	12.56	15.91	10.33	6.65	7.51	10.35
Max	9.44	15.10	15.35	16.66	20.40	16.15	9.79	12.08	13.50
Mean	8.49	14.09	13.01	14.12	18.62	12.99	8.26	10.45	11.86
Std. Dev.	0.68	0.74	1.83	1.39	1.41	1.78	1.04	1.45	1.07
Mean p, q	8.40	13.68	12.44	14.58	19.25	12.11	8.19	9.90	11.30
Mean p_2, q_2	8.60	14.56	13.67	13.60	17.89	14.00	8.34	11.03	12.49
1996–2001									
Min	6.32	9.33	8.68	6.16	9.75	8.79	5.10	5.95	8.91
Max	8.66	10.35	13.83	13.00	16.63	13.41	7.96	9.58	11.22
Mean	7.63	9.87	11.58	9.05	13.36	10.72	6.42	8.44	9.89
Std. Dev.	0.73	0.33	1.52	2.13	2.12	1.59	0.91	1.25	0.79
Mean p, q	7.71	9.89	11.14	9.80	14.69	9.99	6.51	7.93	9.50
Mean p_2, q_2	7.54	9.85	12.08	8.19	11.85	11.56	6.32	9.02	10.34
2000–2001									
Min	6.97	4.48	6.88	1.54	7.21	4.46	2.84	4.24	6.45
Max	8.66	10.38	13.83	13.06	16.63	13.41	7.96	9.86	11.22
Mean	8.70	6.33	9.06	5.29	12.53	7.78	4.61	7.15	7.92
Std. Dev.	0.78	0.84	1.20	2.54	3.17	1.96	1.01	1.57	0.93
Mean p, q	8.84	6.73	8.68	5.99	14.86	6.62	4.89	6.26	7.34
Mean p_2, q_2	8.53	5.86	9.49	4.48	9.87	9.11	4.30	8.17	8.57

Table 33: Balassa-Samuelson inflation simulations under the assumption Δp^* equals 2% and with the inflation differentials in tradables set at the historical values. The values for the other variables are at the average values for the periods specified, except for real per capita GDP and real total consumption growth, which are taken from Wagner and Hlouskova (2004). The three panels correspond to the periods over which the average values for the explanatory variables (except for per capita GDP and total consumption) are taken.
Min, *Max*, *Mean* and *Std.Dev.* denote the minimum, maximum, mean and standard deviation of the implied inflation rates for all 15 equations. *Mean* p, q and *Mean* p_2, q_2 denote the mean over the corresponding sub-groups of equations only.

Appendix C: Description of Implemented Bootstrap Algorithms

Bootstrapping the panel unit root and panel cointegration tests used in this paper requires to take two issues into consideration. The first is non-stationarity of certain quantities (all tests applied have the null of a unit root in the panel, and correspondingly of no cointegration). The second issue is the serial correlation allowed for in the innovation processes.

Both issues can be handled by resorting to appropriate bootstrap procedures. Bootstrap procedures for non-stationary processes are in the meantime relatively well understood, see e.g. Paparoditis and Politis (2003). In our application we have to take into account in addition the extremely small time dimension of our panels. For this reason, one part of our bootstrap procedures fits an autoregression to the residuals of the unit root test equation respectively of the cointegrating regression. Bootstrapping is then based on the residuals from these autoregressive approximations, which should resemble white noise. For our case with $T = 9$ this might be preferable to some block-bootstrap procedure. For comparison, however, we have also implemented the so called *residual based block* bootstrap (RBB) procedure of Paparoditis and Politis (2003), which has certain asymptotical advantages in terms of power compared to the other procedures implemented, compare Paparoditis and Politis (2002).

Since we are in a panel situation, we can also think about bootstrap procedures that preserve some cross-sectional correlation patterns that may be present. A simple way of doing this is to re-sample residuals according to the same re-sampling scheme for all units. Note, however, that none of the tests for unit roots or cointegration applied is designed to allow for correlation across the units. Panel unit root and cointegration tests that allow for correlation across the individual units and that resort to bootstrapping inference are currently investigated i.a. by Chang (2000), Chang (2004) or Chang and Song (2002).

Note that the panel unit root tests and panel cointegration tests are implemented for two different specifications concerning the deterministic components. One, where only (individual specific) intercepts are contained in the test equation respectively the cointegrating regression and the other where both intercepts and trends are contained. We only discuss the second case in this appendix, the other case follows trivially.

Let us now discuss the bootstrapping algorithms implemented for the panel unit root tests and let us start with the autoregression based algorithms. Denote with $y_{it} \in \mathbb{R}$ the panel data observed for $i = 1, \ldots, N$ and $t = 1, \ldots, T$. Then for each unit the following equation is estimated by OLS:

$$\Delta y_{it} = \gamma_{i0} + \sum_{j=1}^{p_i} \gamma_{ij} \Delta y_{it-j} + u_{it} \tag{18}$$

with Δ denoting the first difference operator (defined on $\mathbb{N}$ here). The lag lengths p_i are allowed to vary across the individual units in order to whiten the residuals u_{it}. Denote with $\hat{u}_{it}$ the residuals of equation (18). Then the following two bootstrap procedures are based on the autoregression residuals.

(i) Parametric: The bootstrap residuals are given by $u_{it}^* = \hat{\sigma}_i \varepsilon_{it}$, where $\hat{\sigma}_i^2$ denotes the estimated variance of $\hat{u}_{it}$ and $\varepsilon_{it} \sim N(0,1)$.

(ii) Non-parametric:[22] Denote with $\hat{u}_t = \begin{bmatrix} \hat{u}_{1t}, & \ldots, & \hat{u}_{Nt} \end{bmatrix}'$ and generate the bootstrap residuals u_t^* by re-sampling $\hat{u}_t, t = p+2, \ldots, T$ with replacement. By re-sampling the whole vector, any contemporaneous correlation across units is preserved in the bootstrap series.

Given u_{it}^* the bootstrap data themselves are generated from

$$y_{it}^* = \begin{cases} y_{it} & t = 1, \ldots, p_i + 1 \\ \hat{\gamma}_{i0} + y_{it-1}^* + \sum_{j=1}^{p_i} \hat{\gamma}_{ij} \Delta y_{it-j}^* + u_{it}^* & t = p_i + 2, \ldots, T \end{cases} \tag{19}$$

As indicated above Paparoditis and Politis (2003) propose a different bootstrap algorithm, the RBB bootstrap, based on *unrestricted* residuals. By unrestricted residuals we mean residuals which are not

[22] For notational simplicity we assume $p_i = p$ for all units here in the discussion.

generated from an equation like (18) where a unit root is imposed, due to estimation in first differences, but from an unrestricted first order autocorrelation. Higher order serial correlation is not dealt with by fitting an autoregression, but by bootstrapping blocks, with the block-length increasing with sample size at a sufficient rate.[23] The implementation of the RBB bootstrap is as follows:

(i) Estimate the equation $y_{it} = \gamma_{i0} + \rho_i y_{it-1} + u_{it}$ by OLS (for each unit).

(ii) Calculate the centered residuals

$$\tilde{u}_{it} = (y_{it} - \hat{\rho}_i y_{it-1}) - \frac{1}{T-1}\sum_{\tau=2}^{T}(y_{i\tau} - \hat{\rho}_i y_{i\tau-1}).$$

(iii) Choose the block-length b and draw $j_0, \ldots, j_{k-1}$ from the discrete uniform distribution over the set $\{1, \ldots, T-b\}$ with $k = \lfloor \frac{T-1}{b} \rfloor$. Here $\lfloor x \rfloor$ denotes the integer part of x. By taking the same realizations j_m for all cross-sections, the contemporaneous cross-sectional correlation is preserved in the bootstrap data.

(iv) Denoting with $m = \lfloor \frac{t-2}{b} \rfloor$ and with $s = t - mb - 1$, the bootstrap data are given by:

$$y^*_{it} = \begin{cases} y_{i1} & t = 1 \\ \hat{\gamma}_{i0} + y^*_{it-1} + \tilde{u}_{ij_m+s} & t = 2, \ldots, kb+1 \end{cases} \tag{20}$$

Note again for completeness that for the tests that only allow for an intercept in the test equation $\hat{\gamma}_{i0}$ above is replaced by zero.

For the panel cointegration tests used in this study we also apply three bootstrap algorithms. These are essentially multivariate extensions of the above. The starting point for the autoregression based bootstrap procedures is now given by

$$y_{it} = \alpha_i + \delta_i t + X'_{it}\beta_i + u_{it} \tag{21}$$

$$X_{it} = A_i + X_{it-1} + \varepsilon_{it} \tag{22}$$

for $i = 1, \ldots, N$, $t = 1, \ldots, T$. Now $\alpha_i, \delta_i \in \mathbb{R}$, $X_{it} = [x_{it1}, \ldots, x_{itk}]'$ and $A_i, \beta_i \in \mathbb{R}^k$. Note for completeness that for the test proposed by Kao (1999) $\beta_i = \beta$ holds for all units. Under the null hypothesis of no cointegration between y_{it} and X_{it} it follows that u_{it} is integrated and that ε_{it} is stationary.

We estimate[24] the above equations (21) and (22) to obtain the estimated residuals $\hat{v}_{it} = [\hat{u}_{it}, \hat{\varepsilon}'_{it}]'$ from

$$\begin{aligned} \hat{u}_{it} &= y_{it} - \hat{\alpha}_i - \hat{\delta}_i t - X'_{it}\hat{\beta}_i \\ \hat{\varepsilon}_{it} &= \Delta X_{it} - \hat{A}_i \end{aligned}$$

Under the null hypothesis $v_{it} \in \mathbb{R}^{k+1}$ is a process whose first coordinate is integrated and whose other coordinates are stationary. These known restrictions can be incorporated into the autoregressive modelling to obtain white residuals by fitting a vector error correction model which incorporates the exact knowledge about the cointegrating space. This is achieved by estimating:

$$\hat{v}_{it} = B_i \hat{\varepsilon}_{it-1} + \sum_{j=1}^{p_i} \Gamma_j \Delta \hat{v}_{it-j} + \mu_{it} \tag{23}$$

with $B_i \in \mathbb{R}^{k+1 \times k}$. The residuals from equation (23), $\hat{\mu}_{it}$ say, should resemble white noise due to appropriate choice of the lag lengths p_i.

As in the univariate case for the panel unit root tests, two bootstrap versions are implemented based on $\hat{\mu}_{it}$.

[23]For an autoregression based implementation of this idea of using unrestricted residuals see Paparoditis and Politis (2002).

[24]Estimation proceeds by unit specific OLS estimation, except for the method of Kao (1999), which rests upon the LSDV estimator to obtain an estimate $\hat{\beta}$ identical across units.

(i) Parametric: Estimate the variance-covariance matrix of $\hat{\mu}_{it}$, $\hat{\Sigma}_i$ say. Denote its lower triangular Cholesky factor by $\hat{L}_i$ and generate the bootstrap residuals $\mu^*_{it} = \hat{L}_i \eta_{it}$ with $\eta_{it} \sim N(0, I_{k+1})$.

(ii) Non-parametric: μ^*_{it} is given by re-sampling $\hat{\mu}_{it}$. By choosing the same re-sampling scheme for all cross-sectional units, the contemporaneous correlation structure is preserved.

The bootstrap series y^*_{it} and X^*_{it} are generated by first inserting μ^*_{it} in (23) and by then inserting the resulting v^*_{it} in (21) and (22).

The multivariate implementation of the RBB bootstrap is based on an unrestricted VAR(1) for $Z_{it} = [y_{it}, X'_{it}]'$ as follows.

(i) Estimate the first order VAR $Z_{it} = A_{i0} + A_{i1} Z_{it-1} + v_{it}$.

(ii) Compute the centered residuals

$$\tilde{v}_{it} = (Z_{it} - \hat{A}_{i1} Z_{it-1}) - \frac{1}{T-1} \sum_{\tau=2}^{T} (Z_{i\tau} - \hat{A}_{i1} Z_{i\tau-1}).$$

Choose the block-length b and draw $j_0, \ldots, j_{k-1}$ from the discrete uniform distribution over the set $\{1, \ldots, T-b\}$ with $k = \lfloor \frac{T-1}{b} \rfloor$ and $\lfloor x \rfloor$ denotes the integer part of x. By taking the same realizations j_m for all cross-sections, the contemporaneous cross-sectional correlation is preserved in the bootstrap data.

(iv) Denoting with $m = \lfloor \frac{t-2}{b} \rfloor$ and with $s = t - mb - 1$, the bootstrap data are given by:

$$Z^*_{it} = \begin{cases} Z_{i1} & t = 1 \\ \hat{A}_{i0} + Z^*_{it-1} + \tilde{v}_{ij_m+s} & t = 2, \ldots, kb+1 \end{cases} \tag{24}$$

Note again for completeness that for the tests that only allow for an intercept in the test equation $\hat{A}_{i0}$ above is replaced by zero.